Explaining the Evolution of Pension Structure and Job Tenure

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Abstract

Average and expected job tenure of workers has fallen significantly over the last two decades. Workers have also experienced a major shift in pension coverage. Traditional defined benefit pensions, designed to reward long tenure, have become steadily less common, while defined contribution pensions, which are largely portable, have spread. The link between job tenure and pension trends has not been closely examined, but it offers insights about both phenomena. This paper uses a contract-theoretic matching model with moral hazard to explain changes in both pension structure and job tenure; we discuss how a richer model with job-specific human capital subject to technology shocks would yield similar results. In our model, a decline in the value of existing jobs relative to new jobs reduces expected match duration and thus the appeal of DB pensions. We argue that these trends are linked to changes in the nature of new technologies. This explanation is consistent with observed trends in technological change, tenure, and pension structure. Our results suggest an additional consequence of technological progress that has not been closely studied.

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1 Introduction

The notion that job stability has declined has become popular in media coverage over the last decade, even during the economic boom of the 1990s. The New York Times, for instance, suggested that “the notion of lifetime employment has come to seem as dated as soda jerks, or tail fins” (Kolbert and Clymer 1996). While there is incomplete agreement over the extent of this trend, most data sets show that job tenure, especially of male workers, has fallen over the last two decades. Average tenure of male full-time employees in the Survey of Consumer Finances fell almost 10%, from 9.7 to 8.8 years, and expected remaining job tenure dropped more. Average tenure of female full-time employees rose and then fell, suggesting that their rising attachment to the labor force was tempered by an overall decline in tenure.

Workers have also experienced a major shift in pension coverage since the early 1980s. Traditional defined benefit pensions, designed to reward long tenure, have become steadily less common, while defined contribution pensions, which are largely portable, have spread. The link between job tenure and pension trends has not been closely examined, but it offers insights about both phenomena.

Analyzing this link allows us to bridge some key gaps in the literatures on job stability and on the structure of compensation. First, we develop a matching model with endogenous job destruction that can explain the use of deferred compensation contracts and their connection to job duration. Earlier models of pensions typically did not incorporate uncertainty, nor make explicit the nature of the worker’s outside option – both of which crucially affect the value of deferred compensation contracts. Earlier models of job matching rarely incorporated the use of deferred compensation. A group of recent search papers has begun to analyze tenure-based contracts designed to deter on-the-job search; this paper uses a model with a simpler form of moral hazard to highlight how changes in the economic environment alter the feasibility of such contracts.

Second, we discuss what kinds of shifts in the stochastic productivity process can explain the observed trends in job tenure and pension structure. The model does not require a change in the productivity of new matches. Instead, we focus on two less drastic possibilities: (i) an increased frequency of shocks that reduce the value of existing matches relative to new matches or (ii) an increase in uncertainty. Thus, the model provides possible explanations for the observed decline in job tenure, which few researchers have analyzed. It also offers a new, endogenous explanation for the decline in DB pensions that differs from the focus of other researchers on exogenous changes in pension regulation.

Third, we argue that new technologies have reduced the value of existing jobs relative to new jobs and, perhaps, raised uncertainty. We also demonstrate that observed trends in technological change, tenure, and pension structure support the empirical implications of the model. Many other studies have shown that new

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1 We have appropriated this quote, with thanks, from Neumark, Polsky, and Hansen (1999).
2 Burdett and Coles (2003), Stevens (forthcoming), and Friedberg, Owyang, and Sinclair (2003).
technologies like computers require new skills, replace tasks done previously by unskilled workers, increase the complexity required in many remaining tasks, and induce a reorganization of work within a firm.

The paper is organized as follows. In Section 2, we discuss pensions and pension regulation. We argue that regulatory changes do not fully explain the shift in pension structure. If, nevertheless, regulatory changes have had an effect, our model provides a framework to understand the welfare implications of these interventions.

In Section 3, we review past research on the function of DB pensions. In a series of papers, Lazear argued that DB pensions are designed to encourage longer tenure in the face of imperfect monitoring of worker effort. We discuss other reasons why longer tenure may be valuable, such as motivating match-specific investment.

In Section 4, we develop a matching model and incorporate Lazear’s notions of DB pensions. We show that a contract that defers compensation elicits optimal effort and longer tenure. The deferred payment is conditioned on tenure, mimicking a DB pension. However, the contract may break down in the face of shocks to the output process which make it riskier to get bound into a long-term relationship. Lastly, we allude to reasons why DC pensions may be used when DB pensions are abandoned.

In Section 5, we use our model for a different purpose. Instead of generating an endogenous shift in pension structure, we analyze the efficiency consequences of government restrictions on the use of pensions.

In Section 6, we discuss the empirical predictions generated by the model. A lack of data makes it difficult to estimate our model directly. Instead, we show that trends in pension structure, job tenure, and technological change are consistent with our explanation for the evolution of pension structure and job tenure. We find that job tenure is related to pension structure, that the value of long-term jobs appears to have dropped, and that greater declines in job tenure occurred in industries with higher rates of technological progress.

In Section 7, we conclude by linking our results to other research on the nature of new technologies. Many of the phenomena identified in earlier studies support our explanation about why the value of long-term jobs has declined. Our results suggest a further consequence of technological progress that has not been closely studied.

2 Background

In this section, we set the stage by discussing trends in job tenure and pension structure. We also discuss the structure of typical DB and DC pensions. Lastly, we contend that, while pension regulation has changed a great deal, it does not fully explain the observed trends in pension structure.

3Similar implications arise in a richer model with match-specific capital that gradually decays in the absence of new investment and is vulnerable to technological shocks (Engemann, Friedberg, and Owyang 2003). Friedberg, Owyang, and Sinclair 2003 showed that DB pensions may also be used to deter on-the-job search by workers.
2.1 Trends in job tenure and pension structure

We show that both current and expected remaining job tenure fell, so total expected job duration fell by roughly 6-16%, depending on the time period and sample. Over the same period, DB pension coverage declined a great deal. Our theoretical and empirical analyses later on will show the links between these trends.

To examine these trends, we use data on individuals from the Survey of Consumer Finances (SCF). The SCF includes information on actual and expected job tenure and on pension coverage. It began in 1983 and surveyed a new cross-section every three years, offering the longest consistent information on pension coverage among individual-level data sets. Other data sets do not report expected tenure nor include information on both tenure and pension structure. The primary disadvantages of the SCF are relatively small sample sizes (roughly 5000 individuals per survey) and very aggregated information on industry and occupation (only 6-7 classification codes reported).

2.1.1 Trends in job tenure

Current job tenure. Average job tenure of male full-time employees in the SCF fell from 9.7 in 1983 to 8.8 years in 1998. Average tenure of female full-time employees rose from 7.4 years in 1983 to 8.1 years in 1992 and then fell back to 7.3 in 1998.

Table 1 about here

Table 1 shows average job tenure broken down by gender, education, and years in the labor market (as measured by potential experience since completion of education). Average tenure of men with 0-5 years of potential experience – those least likely to have DB pensions – declined from 3.0 years in 1983 to 1.8 years in 1998. Average tenure of those with 6-15 and 16-25 years of potential experience declined from 5.1 to 4.5 years and from 10.0 to 8.7 years, respectively. The lower panel of Table 1 shows that tenure fell for workers who attended college as well as those who did not.\footnote{We omit data from the 1986 SCF, which asked fewer questions about pensions and only surveyed respondents from the 1983 SCF.}

Changes in job tenure among women apparently reflect a combination of cohort-specific increases in labor force attachment and secular declines in job tenure. For example, tenure rose and then fell a little for those 16 or more years of potential experience, while it tended to remain steady early on and then fell more for those with

\footnote{The CPS has pension supplements, but the last one was in 1993, and the wording of questions changed over time. The Survey of Income and Program Participation has job tenure and mobility data but weak information about pensions. The National Longitudinal Survey of Youth offers job histories but for a limited age range. The Health and Retirement Study offers job histories and very detailed information about pensions, but for a limited age range and only beginning in 1992. The pension data on Form 5500, reported to the government by employers, lacks information on worker and job characteristics.}

\footnote{Thus, declines in tenure are not limited to less-skilled workers who experienced other negative labor market trends over this period.}
less potential experience.

*Expected remaining job tenure.* The SCF also asked respondents how long they expected to continue working for their current employer – providing a direct measure of expected job duration, which is a key element of the model we present later. Expected tenure data are noisier than actual tenure data but show a clear decline as well, so the decline shown in Table 1 reflects more than a one-time reshuffling of workers into new job with unchanged levels of expected tenure.

Among full-time employees aged 21-59, expected remaining tenure for men fell from 18.6 in 1983 and 16.2 in 1989 to 14.7 in 1998 and for women fell from 15.9 in 1983 to 12.5 in 1989, rose to 14.4 in 1992, and then fell to 12.8 in 1998. Table 2 shows expected remaining job tenure by gender, education, and years of current tenure. Among men, expected remaining tenure fell most for those with shorter current tenure – which confirms that the expected duration of new matches has declined. For example, for those with 0-5 years of current tenure, expected remaining tenure fell from 19.6 years in 1983 and 16.7 in 1989 to 14.5 in 1998. It fell by less for men with current tenure of 6-15 and 16-25 years, while it changed little for men with more than 25 years of tenure. The lower panel of Table 2 shows that expected tenure of more-educated men fell more uniformly than expected tenure of less-educated men. This appears consistent with our hypothesis about the impact of skill-specific technological change, which should affect skilled workers disproportionately if it is eroding existing skills and generating uncertainty about future required skills.

Table 2 about here

Again, changes in expected job tenure among women reflect a combination of rising labor supply together with declining job tenure. As with men, expected remaining job tenure for more educated women fell relative to less educated women.

*Total expected job tenure.* Adding together current tenure with expected remaining tenure yields an estimate of total expected job duration. For men, total expected tenure fell from 27.3 years in 1983 to 24.6 years in 1989 and 23.0 years in 1998, a 15.5% decline between 1983 and 1998 and a 6.3% decline between 1989 and 1998. For women, the total went from 22.8 years in 1983 and 22.1 years in 1992 to 19.9 years in 1998. Thus, expected tenure fell by 12.9% between 1983 and 1998 and by 10.3% between its peak in 1992 and 1998.

*Other research on trends in job tenure.* Early researchers, spurred by anecdotal reports of a decline in long-term jobs, argued that job tenure appeared to be stable (Diebold, Neumark, and Polsky 1996, 1997; Farber 1996). Since then, researchers have found mounting evidence of a decline in male job stability in the Current

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7The wording and organization of survey questions regarding future work plans was different enough that the 1983 SCF may not be comparable to later years. Therefore, we discuss how expected tenure changed both since 1983 and since 1989. We restrict the age range of this sample to 21-59 in order to abstract away from any changes in educational attainment or retirement.
Population Survey (Neumark, Polsky, and Hansen 1999; BLS 2000), the Panel Study of Income Dynamics (Jaeger and Stevens 1999), and the National Longitudinal Surveys of young cohorts (Bernhardt, et. al., 1999), although not the Survey of Income and Program Participation (Gottschalk and Moffitt 1999; Bansak and Raphael 1998). Farber (1997) noted that the decline in tenure in the CPS was not matched by an increase in layoffs, indicating that the increases in mobility have been to some extent voluntary. None of these earlier papers used the SCF to investigate changes in job stability.

2.1.2 Trends in pension structure

Figure 1 shows that DB pensions have become much less common and DC pensions have become more common. Among full-time employees with a pension, 69% had a defined benefit (DB) plan and 45% had a defined contribution (DC) plan in 1983, while 40% had a DB plan and 80% had a DC plan in 1998. Overall pension coverage declined somewhat at the same time, from 67% of full-time employees in 1983 to 58% in 1998. Later, we show that remaining DB pensions seem to have declined in value as well.

A more detailed analysis of pension trends suggests a relationship to underlying structural changes in the economy. Pension coverage did not shift uniformly in all types of jobs, but rather that at varying rates by industry, occupation, and education level — in other words, by type of job and type of worker. In Table 3 we use analysis-of-variance to decompose pension trends by type of job and worker. Year main effects explain just under half (48%) of the over-time variation in DB pension coverage — so half of the decline occurred uniformly across types of jobs. Year-industry interactions explain 22%, indicating different changes in DB pension coverage across different industries; year-occupation interactions explain 13%, and year-education interactions explain 15%. When we perform the same analysis of DC pension coverage, year main effects explain 34% of the variation across years; the other important factors are year-industry interactions (explaining 32% of the time variation) and year-occupation interactions (explaining 25%). Friedberg and Owyang (2002b) offered detailed information about pension trends within industries and occupations.

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9 In the SCF, little of the aggregate trend can be explained by the movement of workers across broad industries and occupations; within-occupation and within-industry changes in pension structure dominate. As discussed later, research using detailed industry and occupation codes reported in Form 5500 data indicates that up to half of the shift in pension coverage arises from workers moving from jobs that typically offer DB pensions to jobs that typically offer DC pensions.
2.2 The structure of pensions

Defined benefit pensions. DB pensions offer a defined payout to workers after they leave an employer. While the specific parameters of DB pensions vary a great deal across employers, there are a few common elements. The benefit is typically an annuity that is proportional to either the worker’s average or final salary, with the proportion increasing, sometimes nonlinearly, in tenure. We can summarize the flow in terms of pension wealth $P_t$, defined as the actuarially discounted real present value of the worker’s expected future pension benefits, if the job ends at year $t$. Pension wealth accrual is the discounted change in pension wealth $\frac{1}{1+r}P_{t+1} - P_t$, if the employee works one additional year and then leaves.

Pension wealth can reach quite high levels by the end of one’s career. Allen, Clark, and McDermed (1988) estimated that the pension loss associated with switching jobs for the average worker aged 35-54 is approximately half a year’s earnings. Among older workers with a pension in the Health and Retirement Study (HRS) in 1992, median pension wealth is around $200,000, assuming that workers eventually leave their job at age 65.

Just as importantly, the path of DB pension wealth accrual is typically characterized by sharp spikes. Figure 2 shows pension wealth accrual in a representative DB pension plan from the 1992 HRS. The first spike, worth about $60,000, occurs when the worker vests, or becomes eligible for future benefits. Maximum vesting dates of 10-15 years were established in 1974 and have since been lowered to 5-7 years. While vesting yields a claim to future benefits, pension wealth continues to rise as the future benefit increases with earnings growth, tenure, and the approach of retirement.

Another spike worth over $100,000 occurs when the worker reaches the plan’s early retirement date (ERD), often at ages 55-60 with at least 20 years on the job. At the ERD a retiree can first begin to receive cash benefits. The spike in Figure 2 results from an additional discrete jump in the pension benefit at the ERD. The early benefit is generally less than the full benefit available at the normal retirement date (NRD), and if it is significantly less, then another spike in pension wealth occurs at the NRD. Frequently, though, the penalty for retiring early is small, as is the case in Figure 2. At some point, pension accruals swing around and turn negative after the ERD or NRD, since the worker foregoes income by not retiring.

To sum up, DB plans typically discourage worker mobility for many years after a worker starts a job and later encourage retirement when pension accruals turn negative. Gustman and Steinmeier (1995) pointed out, nevertheless, that pension wealth may be quite small at the start of a job. Therefore, while the reward to staying

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10 The HRS sample was aged 51-61 in 1992. The HRS obtained detailed information about pension plans from employers of survey respondents; the median value of pension wealth was computed for those plans for which employers reported information. The pension plans shown in Figure 2 have been slightly altered, as described in Friedberg and Webb (2000, 2003), to protect confidentiality.
in the job and accruing DB pension wealth is high at that point, so might be the gain to changing jobs, if it entails a pay raise worth a few percentage points and the same type of compensation path afterwards. They argued, therefore, that the primary effect of DB pensions is to deter mobility of longer-tenure workers, rather than new workers. When we present our model later, we will discuss extensions involving investment in match-specific capital which are laid out in Engemann, Friedberg, and Owyang (2003) and which capture the distinction between mobility incentives of shorter and longer-tenure workers.

**Defined contribution pensions.** Employer contributions and mandatory employee contributions to DC plans are a form of deferred compensation, since employees are typically forbidden from withdrawing or borrowing against their plan balances (Mitchell 1999). The accrual of DC pension wealth is simple: contributions go into an account which earns a return, and the account is portable after vesting. Thus, DC pension wealth \( P_{DC}^t \) after vesting is \( P_{DC}^{t-1}(1 + r_t) + c_t \), where \( r_t \) is the rate of return on last period’s assets and \( c_t \) is this period’s contribution. Most DC pensions have vesting periods that are either immediate or less than two years (Mitchell 1999).

The smooth path of DC pension wealth accrual shown in Figure 2 stands in stark contrast to the path of DB accrual. Thus, DC pensions are tenure-neutral after vesting, in contrast to DB pensions. Friedberg and Owyang (2002a) discuss the differences between DB and DC plans in more detail.

In addition, cash balance plans are hybrid pensions that are becoming more common. Cash balance plans accrue pension wealth like DC plans but are technically DB plans in terms of funding standards. For our purposes, the funding mechanism is irrelevant, and we would classify these as DC plans because pension wealth accrues smoothly. Thus, they constitute a new element in the trend away from tenure-based accruals in pension wealth.

2.3 Regulation of pension plans

**Summary of regulatory changes.** Employer-provided pension plans were largely unregulated until the Employee Retirement Income Security Act (ERISA) of 1974. Since then, the government has frequently altered and tightened pension regulations. These changes have focused on three aspects of pension provision:

- Funding of DB pensions. ERISA established the Pension Benefit Guaranty Corporation, which insures DB pensions. ERISA and subsequent laws also set increasingly strict limits on both under- and over-funding of pensions.

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11 Voluntary contributions by employees do not strictly constitute deferred compensation but confer tax benefits that are not available for all forms of saving. The tax treatment of DB and DC plans is similar — contributions are tax-deductible, returns accumulate tax-free, and both DB pension benefits and DC pension withdrawals are taxable income.
12 19% of Fortune 1000 firms sponsored cash balance plans in 1999, and the majority were less than five years old (United States General Accounting Office 2000).
13 There is little concern about misclassifying cash balance plans in our data analysis because they were still quite uncommon at the end of our SCF sample period in 1998.
of DB pensions.

- Tax incentives for DC pensions. Post-ERISA legislation established tax deductions for new varieties of DC pensions, particularly those involving voluntary contributions to 401(k) plans.

- Constraints on the structure of both types of pensions. One goal of new regulatory constraints on pension structure has been to strengthen rules ensuring that employers do not design plans in order to favor high-earning employees. Another goal has been to reduce losses in pension wealth experienced by workers who leave a job early. For example, ERISA established a maximum vesting period of 10-15 years, shortened to 5-7 years by subsequent legislation. A third goal has been to ensure that workers who stay in a job for a long time continue to accrue pension benefits. Thus, a 1986 law prohibited age-based cutoffs on the accumulation of DB pension benefits.

Impact on pension structure. While enhanced tax incentives can explain why DC pensions have spread, it cannot explain why DB pensions have become less common, since a worker can have both types of plans. Researchers have suggested at least three other ways in which regulatory changes may have caused the shift away from DB pensions. However, we argue that none of these fully explain observed trends in pension structure, and that some changes may have had the opposite effect.

First, as pensions have been increasingly regulated, the costs of administering both DB and DC plans have risen. Ippolito (1995) reported estimates from the Hay-Huggins Company (1990) indicating that only very small DB plans grew relatively more expensive compared to very small 401(k) plans, while average administrative costs of medium-sized and large plans rose at similar rates. Using longitudinal data on pension plans, Kruse (1995) concluded that rising administrative costs might explain some but not all of the decline in DB pensions during 1980-86.

Second, Clark and McDermed (1990) argued that many of the regulatory changes limited the extent to which DB plans could be designed as incentive contracts of the type we model later. For example, before ERISA many DB pensions only vested at the normal retirement date (Ippolito 1988). However, DB pension wealth can still accrue highly nonlinearly. The DB plan in Figure 2 delivers a much greater fraction after twenty years of tenure than when it vests, and it begins to lose value shortly after the early retirement date in spite of a prohibition on age-related cutoffs.

Third, Ippolito (2001, 2003) claimed that regulatory changes involving reversion taxes allowed companies to escape their DB pension obligations more easily than before. Once some companies took advantage of this, he claimed, it undermined the confidence of other workers in the implicit pension contract. However, a fourth effect related to confidence may be important as well. Some of the regulatory changes involving pension funding
enhanced the appeal of DB plans to workers. In the model developed in this paper, we disregard the possibility that an employer could renege on the DB pension – for example, by underfunding pensions and then declaring bankruptcy. If reneging is a concern, however, then new funding standards and pension insurance will increase the willingness of workers to accept DB pensions.

2.4 Past research on the shift in pension structure

Past research suggests that the evolution of pension structure has been associated with structural shifts in the economy. That is a major reason why we argue that regulatory changes fail to completely explain pension trends. A series of papers showed that workers have shifted from jobs that typically offer DB plans to jobs that typically offer DC plans. These papers, including Clark and McDermid (1990), Gustman and Steinmeier (1992), Kruse (1995), Ippolito (1995), and Papke (1999), used Form 5500 data reported by employers to investigate changes in pension structure. While each of the papers showed that DB pensions remain more prevalent in large firms, industries such as manufacturing, and unionized jobs, several also suggested that the proportion of workers in such jobs declined. According to Gustman and Steinmeier (1992) and Ippolito (1995), for example, the movement of workers across jobs might explain half of the shift in aggregate pension structure.

Other papers have shown that inequality in pension coverage has grown along with earnings inequality. Using data on individuals from the Current Population Survey, Bloom and Freeman (1992) and Even and Macpherson (2000) showed that inequality in pension coverage by skill group has increased, mirroring trends in earnings inequality – a further indicator of structural change related to use of DB pensions.

3 Theories of DB Pensions

Past theoretical research has sought to explain the incentive effects of DB pensions. We build on the idea, originally developed by Lazear and others, that DB plans are designed to encourage optimal effort and longer tenure.

3.1 DB pensions as incentive contracts

DB pension incentives while young. In a series of papers summarized in Lazear (1986), Lazear developed models in which employers structure compensation to deter shirking by workers whose effort cannot be observed perfectly. A DB pension, whose value rises with job tenure, motivates effort by workers who do not want to get fired and

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14 Employers report pension plan data annually to the government on Form 5500. Form 5500 data covers the universe of plans with more than 100 participants, with comparatively little measurement error. Its primary disadvantage is the lack of information about characteristics of covered workers.
lose their “bond”.15

Other reasons to encourage longer tenure. Sunk costs provide another motive for the use of DB pensions. For example, if hiring or firing entails fixed costs, then employers will only hire if expected job duration is long enough. The same factor affect employers’ decisions to train workers and workers’ decisions to invest in job-specific training. In this paper we focus on moral hazard as a motive for DB pensions, while in Engemann, Friedberg, and Owyang (2003) we incorporate job-specific investment and vacancy posting costs. A different motive for tenure-based contracts, though with some similar implications, is offered in Burdett and Coles (2003), Friedberg, Owyang, and Sinclair (2003), and Stevens (forthcoming); employers defer compensation in order to discourage on-the-job search by workers who seek better outside offers.

DB pension incentives while old. Lazear (1983) argued that DB pensions also function as severance pay to encourage efficient retirement when workers age. The pension works in concert with a rising wage profile, another element of the incentive contract. If a fixed amount is to be paid over some duration as wages, the payments can be structured to rise over time, paying a worker less than her marginal product early and more later.16 However, rising wages encourage workers to stay in the firm too long, particularly when there is a retirement motive generated by declining marginal productivity or rising marginal utility of leisure.17 Also, the rising wage profile creates an incentive for employers to violate the implicit contract by firing workers after getting most of the benefits of the increased productivity. Both these problems that undermine the implicit contract may be resolved by a DB pension.18 A DB pension encourages the worker to retire at the “right” age, before pension accruals turn negative. That, in turn, reduces the incentive of employers to fire older workers at an arbitrary old age.

We chose not to incorporate a rising wage profile in our model, since empirical tests of theories of rising wages have not yielded clear conclusions.19 We could extend our model to generate a rising wage profile as in Ippolito (1994), if we imposed restrictions on the extent to which compensation could be deferred through the DB pension.

We also chose not to include an explicit retirement motive, since our model generates an endogenous termination date without it. Incorporating a retirement motive might be important if there were evidence of changes in retirement motives that coincided with the shift in pension structure. This seems unlikely, though, since

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15 Related ideas also appeared in Becker and Stigler (1974).
16 Ippolito (1994) endogenized the rising wage profile. In his model, wage tilt is necessary when the DB pension is too small to deter a worker from quitting after receiving an attractive outside option. Akerlof and Katz (1989) showed that, in the absence of up-front performance bonds, a rising wage profile is insufficient to deter shirking early in the career; a pension can deter shirking and is often cheaper to the firm than allowing shirking.
17 In contrast, Akerlof and Katz (1989) assumed an exogenous retirement date.
18 Research in this vein was prominent following the elimination of mandatory retirement beginning in the 1970s. Alternatively, mandatory retirement gets the worker to retire at the optimal date in Lazear (1979), while the DB pension deters shirking just before the terminal date. However, research by Burkhauser and Quinn (1983) and Stern and Todd (2000) has raised doubts about the empirical effect of mandatory retirement rules.
19 For example, Ippolito (1991) found that wage tilt had no significant effect on job tenure, while DB pensions did. More generally, the literature lacks consensus on the extent to which wages and productivity diverge over the life cycle, and the extent to which such divergence is rooted in moral hazard versus human capital explanations.
pensions of younger and shorter-tenure workers have changed much more than pensions of older workers. Also, if anything, the move away from DB plans may have increased firms’ use of temporary early retirement inducements (Lumsdaine, Stock, and Wise 1990; Brown 2000).

3.2 Other possible motives for DB pensions

Screening. Some researchers have suggested that employers offer DB pensions in order to attract more productive workers, rather than to elicit higher productivity after workers are hired. In Viscusi (1985) and Ippolito (1994), firms gain when workers stay in the same job, but workers have private information about the expected value of future outside offers. Some implications of these screening models are problematic, though. Workers who accept a pension in Ippolito (1994) are those with the lowest expected value of the outside offer – yielding the counterintuitive outcome that workers with the highest productivity in a particular firm are those with the lowest productivity outside of it.

Empirical tests of screening and self-selection motives for DB pensions have been attempted but run into difficulties in resolving identification problems. Allen, Clark, and McDermid (1993) estimated a model of turnover and pension coverage in order to distinguish the importance of various motives for DB pensions. Identification relied on the exclusion of geographic dummies from the pension equation, as well as functional form. They found some role for self-selection but on observables rather than unobservables, which calls into question the need for pensions as a screening device. Using a similar approach, Even and Macpherson (1990) analyzed the role of pensions in sorting women, who have higher exogenous quit rates, into jobs with low rates of pension coverage.\footnote{Their analysis raises an interesting issue – increased labor supply by women, who have shorter average tenure, might explain the shift away from DB pensions. However, our analysis of SCF data shows that industries and occupations with greater gains in the share of women in employment did not experience greater declines in the use of DB pensions.}

Another important point is that an endogenous explanation for the shift in pension structure in this class of models involves a decline in the value of screening. This seems implausible given other labor market trends such as the growth in earnings inequality among workers with similar skills, as measured by education and labor market experience, which has been widely interpreted as an increased return to unobserved ability (Gottschalk 1997). An increase in the return to unobserved ability suggests an increased need to screen workers and thus a greater value of deferring compensation.

In sum, theories related to screening lead to some counterintuitive predictions and fail to offer a ready explanation for the shift in pension coverage. Therefore, we have chosen not to incorporate a screening motive for pensions.

Labor unions. The observed link between unionization and pension coverage has led some researchers to focus on theories related to union bargaining. Freeman (1985) argued that unions give a stronger voice to older,
less mobile workers who gain from pensions at the expense of younger workers with higher quit rates. However, Gustman, Mitchell, and Steinmeier (1994) pointed out that the use of the pension to defer compensation in this context is puzzling in light of the generally flatter age-earnings profile observed in unionized firms. Freeman also suggested that unionization itself may reduce mobility and increase the willingness of workers to accept a pension. If workers already have a low probability of quitting, though, a DB pension is not particularly useful in incentive models of pensions.

Another implication of the model we develop later is important to note – the impact of a decline in a worker’s bargaining power is ambiguous. A decline in bargaining power (associated, say, with a decline in unionization) will make shirking more attractive and thus increase the value of the pension contract, though at the extreme it destroys the pension contract entirely because the contract can no longer deter shirking at all.

Empirical evidence on the link between unionization and pension trends is also ambiguous. Although DB plans are more common in unionized jobs, many non-unionized jobs offer them as well. Among unionized workers with a pension in the SCF, 70% have a DB plan and 53% have a DC plan, and among nonunionized workers, 45% have a DB plan and 75% have a DC plan. Also, while both DB pensions and unionized jobs have disappeared over time, the evidence presented above and cited in earlier research suggests that the shift out of unionized jobs does not explain a great deal of the change in pension structure.

4 A Model of Pensions

We develop an incomplete-contracting job-matching model that incorporates insights from past research on DB pensions. Matching models offer a rich representation of the labor market and the effects of uncertainty which is absent from earlier models of pensions. However, search and matching models have tended to focus on the rate and duration of unemployment and feature exogenous job destruction, while ours emphasizes the duration of employment and features endogenous job destruction, which motivates the use of pensions.

In order to develop our arguments, we present a baseline Nash bargaining model (Model N) and then a Nash bargaining model with moral hazard (Model MH), which build on den Haan, Ramey, and Watson (2000, hereafter DRW). As in DRW, moral hazard induces endogenous match destruction. We propose a pension-contract model (Model P) that discourages moral hazard and eliminates inefficient match destruction, although we do not demonstrate that it is the only contract that would do so. We will also mention, where appropriate, extensions of this model which are developed in Engemann, Friedberg, and Owyang (2003) and Friedberg, Owyang, and

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21 Mortensen and Pissarides (1999) offered a survey of matching models.
22 Valletta (1999) discussed how extensions to a similar model by Ramey and Watson (1997) could help explain the decline in job tenure, but he did not write down such a model or focus on pensions, as we do. Ramey and Watson (1997) modeled bilateral shirking in a pure contract-theoretic framework without search. Our approach extends the severance contract which they outlined to the search and matching model employed in DRW.
Sinclair (2003). After presenting the pension model, we discuss changes in the productivity process that would lead agents to abandon the pension contract. The Appendix contains a list of variables used in the model.

4.1 The baseline model with Nash bargaining

The matching market. As in DRW, a continuum of atomistic unemployed workers and firms who are searching in the labor market in a given period meet each other with probability $\lambda$.$^{23}$ The matched worker and firm $i$ get an output draw $Y_{i,t}$ and decide whether to produce. If they do not produce, they return to the matching market next period. They decide to produce if the output draw exceeds a threshold value $R$, reflecting the surplus from producing today and from the option to get another output draw and produce in future periods.

Output. Output $Y_{i,t}$ is drawn from a distribution $F(y_{i,t})$ which is the same for all new matches. Thus, while matches are heterogeneous in their actual production draws, workers and firms are identical ex ante. For simplicity, we assume that the distribution $F(y_{i,t})$ is fixed over time and therefore suppress both the time and match identifying subscripts.$^{24}$ Later we will assume that output is drawn from a standard uniform distribution in order to illustrate the important characteristics of the model.

Production under Nash bargaining. In each period, the agents decide whether to continue producing or whether to draw their outside options and rejoin the labor market. They do this by comparing the match output with the outside options if they break up the match. If they produce, the agents split the match surplus, with a share $\theta$ going to the worker and $1-\theta$ going to the firm.

If the match breaks up, the worker and firm receive $b^w$ and $b^f$ from their contemporaneous outside option and expect $\phi^w$ and $\phi^f$ from re-entering the matching pool. If they continue to produce, the current value of the match is $Y + g(R)$, where $g$ is the continuation value of staying in the match and depends on a threshold output level $R$ satisfying

$$R + g(R) = \phi + b,$$

where $\phi = \phi^w + \phi^f$ and $b = b^w + b^f$. When $Y_t = R$, agents are just indifferent between continuing or breaking up the match.

Then, we can define the joint continuation value of the match as

$$g(R) = \beta \int_R^{\infty} (y + g(R)) dF(y) + \beta \int_0^R (\phi + b) dF(y).$$

The continuation value equals the discounted value of the match next period if output exceeds the threshold value

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$^{23}$Friedberg, Owyang, and Sinclair (2003) allows workers to search while on the job.

$^{24}$In Engemann, Friedberg, and Owyang (2003), we allow the distribution to drift.
\( R \), plus the discounted value of the outside option if output does not exceed \( R \). Joint surplus from the match is defined as the value of the match less the value of re-entering the matching pool, \( Y + g(R) - \phi \) and gets split between the agents according to the worker’s relative bargaining share \( \theta \).

The value of re-entering the matching pool to the worker and firm is given by

\[
\phi^w = \lambda \beta \int_R^\infty (\theta (y + g(R) - \phi - b)) dF(y) + \beta (\phi^w + b^w) \tag{2}
\]

\[
\phi^f = \lambda \beta \int_R^\infty ((1 - \theta) (y + g(R) - \phi - b)) dF(y) + \beta (\phi^f + b^f). \tag{3}
\]

These values depend on the probability of re-matching \( \lambda \) and subsequently drawing a satisfactory level of output (exceeding the threshold \( R \)) or alternatively remaining in the matching pool until the subsequent period. For ease of discussion, we define the following:

\[
J(R) = \int_R^\infty ydF(y).
\]

\( J(R) \) reflects the range of output conditional on the match persisting \((y > R)\), which decreases as \( R \) increases.

**Solution.** We assume that the distribution \( F(y) \) is standard uniform and constant over time.\(^{25} \) The model can then be summarized by the following equations:

\[
R = \phi + b - g \tag{4}
\]

\[
g = \frac{1}{2} \beta (1 - R^2) + \beta (1 - R) g + \beta R (\phi + b)
\]

\[
\phi = \frac{1}{2} \lambda \beta (1 - R^2) + \lambda \beta (1 - R)(g - \phi - b) + \beta (\phi + b).
\]

The solution to (4), denoted by the subscript \( N \), is obtained from:

\[
R_N = \frac{2b - \beta (1 - \lambda) (1 - R^2)}{2k} \tag{5}
\]

\[
g_N = \frac{2\beta bR + \beta (1 - \beta (1 - \lambda)) (1 - R^2)}{2(1 - \beta) k}
\]

\[
\phi_N = \frac{2\beta b(k - \lambda(1 - R)) + \beta \lambda (1 - R^2)}{2(1 - \beta) k}
\]

\(^{25}\) For a generic distribution of \( Y \), the model can be expressed in terms of \( J(R) \) and of the probability \( \Pr[Y < R_N] \) of match dissolution. The assumption \( F_{t+1}(y) > F_t(y) \) yields the upward drift case of DRW, in which no jobs are destroyed in steady state. \( F_{t+1}(y) < F_t(y) \) yields the downward drift case of Caballero and Hammour (1994, 1996), in which jobs are inevitably destroyed.
where \( k = 1 - \beta(1 - \lambda)(1 - R) \).\(^{26}\) A higher value of the output threshold \( R_N \) raises the continuation value \( g_N \) but also raises the value \( \phi_N \) of re-entering the matching pool. \( R_N \) itself is a positive concave function of the contemporaneous outside option \( b \). Additionally, for high values of \( b \), \( R_N \) increases in the discount rate \( \beta \) and decreases in the probability of rematching \( \lambda \).

The wage paid by the employer to the worker each period will be the worker’s portion of the surplus plus his outside option less his portion of the match continuation value. Under Nash bargaining, this is equivalent to the worker’s share of output:

\[
    w_{N,t} = \theta Y_t
\]

To facilitate discussion, we offer the following definition:

**Definition 1** Joint Net Productivity is defined as \( Y_t + g - \phi - b \). A match is Jointly Productive if

\[
    Y_t + g - \phi - b > 0
\]

and Jointly Unproductive if

\[
    Y_t + g - \phi - b < 0.
\]

In the Nash model, all matches are Jointly Productive.\(^{27}\)

### 4.2 The Moral Hazard model

The following model illustrates the inefficiency that results when unobservable effort on the part of the worker affects future match productivity. While we specify a simple form of moral hazard — low effort today destroys the continuation value of the match — we will indicate how it stands in for a richer model, developed in Engemann, Friedberg, and Owyang (2003). In the richer model, a worker decides whether to invest in match-specific capital that keeps the output distribution from drifting down, while skill-specific technological changes erodes the stock of capital.

**Moral hazard.** Consider now the earlier model, augmented with moral hazard which pays off \( x \) to the agents; for simplicity, we will limit consideration to moral hazard by the worker. A worker who shirks gains \( x w \) this

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\(^{26}\) solves the fixed point of the first equation of (5) and then determines \( g \) and \( \phi \). If we allow \( F_{i+1}(y) \neq F_i(y) \), then \( R \) is time-dependent and (5) does not have a closed-form solution. Note that \( R \) is negative for some values of \( b \), implying no exogenous breakups.

\(^{27}\) If match productivity drifts downward, matches will inevitably become Jointly Unproductive as more of the probability mass drifts below the reservation productivity. If \( R \) is time-dependent, it changes the steady-state separation characteristics; see Mortensen and Pissarides (1999) for details.
period but reduces future match productivity and thus the continuation value $g(R)$. We assume that shirking causes $g(R)$ to go to zero, so the match is severed.\footnote{In Engemann, Friedberg, and Owyang (2003) match productivity is a function of job-specific human capital which must be kept current at a cost $x^w$ to the worker. When the worker fails to update her specific human capital, match productivity falls enough to induce the firm to sever the match.}

Production in the presence of moral hazard. With the addition of moral hazard, the model becomes:

$$R = \max\{x, b\} + \phi - g$$

(8)

$$\phi = \frac{1}{2} \lambda \beta (1 - R^2) + \lambda \beta (1 - R)(g - \phi - b) + \beta (\phi + b)$$

$$g = \frac{1}{2} \beta (1 - R^2) + \beta (1 - R)g + \beta R(\phi + b).$$

The difference from the Nash model is the possible addition of the moral hazard premium $x$ to the reservation productivity threshold. If $x$ exceeds the outside benefit $b$, reservation productivity is raised by the difference.

**Definition 2** As above, a match with moral hazard is Jointly Productive if

$$Y_t + g - \phi - b > 0.$$ 

The match is Incentive Compatible if both of the following are satisfied:

$$Y_t - w + g^f > \phi^f + b^f$$

$$w + g^w > x^w + \phi^w + b^w.$$ 

The first inequality states that the firm’s payoff (current period profit $Y - w$ plus continuation value $g^f$) must exceed its total outside option. The second inequality states that the worker’s payoff (wage $w$ plus continuation value $g^w$) must exceed the value of shirking plus the worker’s outside option.

Figure 3 about here

The resulting Joint Productivity Threshold ($Z$) and worker’s Incentive Compatibility (IC) constraint are graphed in Figure 3. The vertical axis shows $y + g(R)$, match output plus the continuation value, while the horizontal axis shows $\phi + b$, the outside option. Note that IC lies a distance of $x^w - b$ above $Z$, since moral hazard imposes additional requirements on current productivity to sustain the match. Matches below $Z$ are Jointly Unproductive.
and are destroyed. Matches above IC are productive enough that the worker chooses high effort. Matches in the region between IC and Z are broken up because workers choose low effort even though the matches are Jointly Productive. Essentially, \(x^w > b\) creates a wedge between efficient and sustainable matches that require extra productivity in order to overcome shirking. These scenarios are summarized in the following proposition.

**Proposition 3** Suppose that no steady-state displacements occur in the model without moral hazard (i.e., that (7) above holds). For any \(x^w > b\) and nondegenerate \(F(y)\) with finite support, the probability of match dissolution in the model with moral hazard is strictly between zero and one.

The proposition implies that even though matches are Jointly Productive, there exists some \(Y\) for any \(x^w\) such that the match is not Incentive Compatible.²⁹

**Solution.** In the solution to (8) that avoids the moral hazard problem, reservation output must satisfy:

\[
R_{MH} = x^w + \frac{\beta(1 - \lambda)\left(2b(1 - R) - (1 - R^2)\right)}{2k}
\]

where, again, \(k = 1 - \beta(1 - \lambda)(1 - R)\). The expressions for \(g_{MH}\) and \(\phi_{MH}\) are the same as in (5), but if \(x^w > b\) it drives up the minimum output \(R_{MH}\) required to sustain the match, changing the resulting values of \(g\) and \(\phi\). The MH wage

\[
w_{MH,t} = \theta Y_t
\]

will exceed the Nash wage in expectation because the reservation productivity that solves the fixed point problem (9) is greater. Thus, the MH wage in a sustainable match compensates the worker for forgoing the moral hazard payment.

**The impact of moral hazard.** Given the model (8), the worker will shirk if the value of not shirking and sustaining the match (the wage plus continuation value) is smaller than the payoff from shirking (the premium \(x^w\) and outside option). A higher \(\theta\), the worker’s bargaining power and consequent share of future match rents, reduces the incentive to shirk. A higher \(\lambda\), the probability of re-matching, raises the value of the outside option and hence the incentive to shirk. A higher \(b\), the contemporaneous outside option, with \(b < x^w\), increases the reservation threshold \(R_{MH}\) but by less than it would increase \(R_N\). This is because match surplus, and therefore the wage and continuation value, continue to be determined by \(b\), but \(R\) is now determined in part by \(x^w\) as well, so \(b\) has a reduced effect.

To understand the magnitude of the efficiency loss in response to some of these parameters, we compare...

²⁹A formal proof of a similar proposition appears in DRW. They show that for output below the reservation threshold, the derivative of \(R\) with respect to \(x\) is strictly positive. Thus, if no steady state dissolutions occur, \(x\) can be raised such that \(R > 0\), so some matches that are dissolved in the MH model would not be dissolved in the N model.
aggregate output in period $t$ in the MH model, $\tilde{Y}_{MH} = \int_{R_{MH}}^{\infty} dF(y) = 1 - R_{MH}$, with aggregate output in the Nash model, $\tilde{Y}_N = \int_{R_N}^{\infty} dF(y) = 1 - R_N$. The productivity loss resulting from shirking is

$$\Lambda = \frac{\tilde{Y}_N - \tilde{Y}_{MH}}{\tilde{Y}_N} = \frac{R_{MH} - R_N}{1 - R_N},$$

which is always non-negative since $R_N < R_{MH} < 1$.

Figure 4 about here

Figure 4 plots the productivity loss $\Lambda$, shown on the vertical axis, as the shirk premium $x^w$ and the outside option $b$ vary, given other reasonable parameter values ($\lambda = 0.3$, $\beta = 0.95$, $\theta = 0.5$). Since output per period lies between 0 and 1, we analyze values of $x^w$ and $b$ that are of the same order of magnitude. As we noted above, the productivity loss increases with $x^w$ and $b$, since they make shirking more attractive. For example, the productivity loss ranges from 0 to 15% for $b = 0.66$ and $x^w$ rising from 0.66 to 0.7, and it reaches as high as 50% when $b$ and $x^w$ exceed 0.9.

Summary. Matches in the Moral Hazard model are vulnerable to incentives that raise payoffs to the worker today but destroy the future value of the match. This generates inefficient outcomes by forcing the dissolution of matches that are Jointly Productive. Next in the Pension model circumstances under which a deferred payment conditioned on reaching a certain tenure will eliminate inefficient match destruction.

4.3 The Pension model

In the previous subsection, we assumed that the worker obtains the shirk premium $x^w$ in lieu of production when effort is low. This gives the worker the incentive to collect the short-run premium and re-enter the matching pool. Here, we show that a deferred payment conditioned on match tenure – structured like a DB pension – can change the worker’s incentives. The contract induces the worker to devote full effort and can be constructed to satisfy the same Joint Productivity condition as the Nash model above, yielding efficient matches. We demonstrate that the pension contract satisfies the Incentive Compatibility condition and avoids inefficiency associated with moral hazard; extension of the model to satisfy Joint Productivity is straightforward.

The pension contract. Suppose that the firm and worker write a contract $\{\mathbf{w}, W, T\}$ with the following elements:

- The worker collects wage $w = \mathbf{w}$ in each period when he is working and $t < T$.
- The worker collects $W(T)$, a lump sum, if he is still employed at time $T$. 

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We will discuss the choice of $W$ and $T$ later. Without loss of generality, we will set $\overline{w} = 0$ for the rest of this discussion.\footnote{Enforcement considerations or risk aversion on the part of workers (as in Burdett and Coles 2003) would affect the actual tradeoff between $\overline{w}$ and $W$.}

**Enforcement.** We assume that the firm is prevented from breaking up the match if $y_t > R_N$ and $t < T$. We also assume that the match breaks up if $y_t < R_N$, even if $t < T$. Thus, the firm pays out $W$ at time $T$ as long as $y_t > R_N$ each period.\footnote{If the match is dissolved before $T$, the worker gets nothing. Workers will nevertheless agree to the contract if the expected value of $W$ is high enough.} The assumption that firms are prohibited from severing productive matches but allowed to sever unproductive matches rests on the observability of $y_t$ and $R_N$. It is crucial, however; if a firm could not break up a match once a contract were agreed to, the worker would have no incentive not to shirk. Therefore, we must appeal to reputation effects or to age discrimination laws which make it more difficult to fire older workers systematically than to lay off workers when output suffers. In any case, empirical evidence indicates that obvious breach of deferred compensation contracts by employers is infrequent.\footnote{Gokhale, Groshen, and Neumark (1995) found mixed evidence that hostile takeovers in eight firms led to reductions in extra-marginal wages of workers, compared to nearby firms that did not experience hostile takeovers. Pontiff, Shleifer, and Weisbach (1990) found relatively small gains from DB plan termination in firms experiencing hostile takeovers. Petersen (1992) found that DB plan termination was more likely in firms with more valuable implicit promises to workers, but that taxes had a greater effect on termination. Cornwell, Dorsey, and Mehrzad (1991) found little evidence of opportunistic dismissal of pensioned workers in a nationally representative survey.}

**The worker’s incentives.** Under the pension contract, the worker’s continuation value $g^w$ depend on the wage contract $\{\overline{w}, W, T\}$. Again assuming $\overline{w} = 0$, at the outset

$$g^w = \beta^T W(T).$$

We need to demonstrate several things about the worker’s incentives in order to prove that the pension contract is feasible. First, if the worker accepts the contract in period 1, the worker will not sever the match later. The continuation value grows in later periods since the value is fixed but the worker discounts it less. Thus, by induction, she will not sever the match in any period $t > 1$ unless the productivity distribution shifts (which we have not allowed for yet) such that the worker’s outside option grows relatively more valuable.

Next, we summarize in the following proposition the worker’s incentive to shirk after accepting the contract, along with the worker’s incentive to accept the contract at the outset :

**Proposition 4** Suppose the worker’s payoff to shirking is $x^w$. Then, the worker will accept the contract $\{\overline{w}, W, T\}$ as long as the shirk premium satisfies $x^w < g^w + \overline{w} - \phi^w$. Specifically for the case $\overline{w} = 0$, if $x^w < g^w - \phi^w$ in each period, then the worker will choose high effort.

Thus, the worker accepts the contract and does not shirk if the shirk premium is smaller than the value of the contract less the value of the outside option. We can check to see under what circumstances the contract in
which the worker is paid $\pi = 0$ until time $T$ and $W$ at time $T$ satisfies the inequality. Substituting (11) for $g^w$ and substituting for the value of $\phi_N^w$ we solved for above, this requires

$$x^w < \beta^{T-t} W(T) - \theta \beta \frac{b(k - \lambda(1 - R_N)) + \lambda J}{1 - \beta} k$$

for all $t = 1, 2, ..., T$. As we mentioned above, the constraint is most to likely bind the lower is $t$. As time passes, the worker gets closer to the pension payoff and is less likely to shirk and risk getting fired.

Note that condition (12) determines the minimum $W$ necessary to provide the worker with the proper incentives, satisfying both Incentive Compatibility and Joint Productivity. The actual choice of $W$ could be modeled as depending on $\theta$, the bargaining weight that determines the split of current-period surplus. A possible choice of $W$ is the one equal in present value to Nash bargaining each period. In that case, $W = \sum_{t=1}^{T} \left( \frac{1-R_N}{\beta} \right)^t [\theta y_{t+|t} - \bar{w}]$, or the expected future discounted value of the match less each period’s wage payment, where $R_N$ is the severance risk, $y_{t+|t} = E[y_{t+|t} | \Omega_t]$, and $\Omega_t$ is the information available when the contact is written.

In order to understand how condition (12) governs feasible values of $W$ and $T$, we analyze the impact of the threshold level of output $R_N$ and then the fundamental parameters that determine $R_N$. (12) identifies the highest sustainable shirk premium $x^w$ for a given conditional output $J = \int_{R}^{\infty} y dF(y) = \frac{1}{2}(1 - R_N^2)$ and severance risk $Pr[Y_t < R_N] = \int_{0}^{R_N} dF(y) = R_N$. These two quantities are in tension as the reservation threshold $R_N$ changes. Higher $R_N$ reduces conditional output $J$ and hence the value of re-entering the matching pool by reducing the likelihood that a productive match is formed; this raises the sustainable shirk premium for a given $W$. However, higher $R_N$ also raises the severance risk, so that staying matched becomes more uncertain; this reduces the sustainable shirk premium. At low $R_N$, the effect on $J$ dominates the effect on severance risk, making the required pension payoff $W$ for a given termination date $T$ relatively insensitive to changes in $R_N$. As $R_N$ increases, the severance risk takes over and small changes in the reservation threshold have increasing effect on the sustainability of the pension contract.

Figure 5 shows how the minimum value of the pension $W$ that satisfies (12) is affected by some of the model’s fundamental parameters. Specifically, it shows how the minimum $W$, expressed as a percentage of the total expected value of the match at time $T$, changes as the shirk premium $x^w$ and the vesting date $T$ change, given other reasonable parameter values ($\lambda = 0.3$, $\beta = 0.95$, $\theta = 0.5$, $b = 0.5$). Again, since output per period lies between 0 and 1, we analyze values of $x^w$ of the same order of magnitude.
It is apparent from Figure 5 that the promise of the future pension that offers the worker a share of all future expected output has a powerful effect in deterring moral hazard. Thus, raising the shirk premium from 0.5 to almost 1 has little effect on the minimum required $W$ for a given $T$. Figure 5 also shows the tradeoff between the term of the pension and its payoff; as noted above, an earlier termination date $T$ allows for a lower payment $W$, given $x^w$. If $x^w$ is 0.7, for example, a termination date of 25 periods requires a pension worth at least 27% of total expected Nash output, while a termination date of 15 periods requires a pension worth 10% of total output.

However, at sufficiently high values of $x^w$ (approaching or exceeding 1, the maximum value of per-period output) the pension contract is no longer viable. The only way to deter shirking is to continue to increase $W$ as $x^w$ rises, but this is only profitable for the firm if it also extends $T$. Extending $T$ raises the risk that the match will be severed before $T$, though. At some point, which is governed by (12), the firm cannot offer a high enough $W$ to get the worker to accept the necessary increase in $T$, even if the worker’s discount rate $\beta$ gets very close to 1.

Summary. The contract $\{w, W, T\}$ will be accepted by both agents and enhances efficiency when $x^w$ satisfies (12). In the next subsection, we discuss how changes in the productivity process affect the pension contracts that can be constructed to prevent moral hazard.

4.4 Expected Tenure and the Productivity Process

The previous subsection demonstrated how the DB pension (the lump-sum payoff $W$ at time $T$) can resolve the inefficiencies resulting from moral hazard. In the model we laid out above, match productivity does not drift, so the continuation value remains constant. The pension contract will also be effective if match productivity drifts upward, boosting the continuation value over time. In this section, we analyze the implications of other specifications of the productivity process – for example, downward drift that reduces the productivity of existing matches relative to new matches, or an increase in uncertainty. Later on, we discuss the corresponding technology shocks which we have in mind. These features reduce expected job tenure and, if severe enough, render the pension contract infeasible, as the worker no longer accepts deferral of payment because the risk of exogenous separation becomes too high.

For purposes of discussion, we define the following:

**Definition 5** A worker’s expected tenure is

$$E(\tau) = \frac{1}{1 - \mathcal{R}}.$$  

*Downward drift in the productivity of existing matches.* Consider (7) in which matches are initially Jointly Productive. Suppose now that output each period is drawn from successively less favorable probability dis-
tributions, in the sense of first-order stochastic dominance, so \( F_{t+1}(y) > F_t(y) \) for all \( t \).\(^33\) This implies a time-dependent continuation value in which \( g_{t+1}(R) < g_t(R) \). The expression in (4) then implies an increasing reservation productivity \( R_{t+1} > R_t \), since conditional output \( J > 0 \) for all \( y \); only a higher draw will make the agents willing to continue the match in the face of worsened long-term prospects. In the type of human capital model we outline in Engemann, Friedberg, and Owyang (2003), these shocks result from the introduction of a new technology which erodes the value of existing skill-specific human capital.

The resulting condition \( R_{t+1} > R_t \) has implications for job tenure. The severance risk \( \Pr[Y_t < R_N] = \int_0^{R_t} dF_t(y) \) increases when either the distribution gets less favorable or reservation output rises, increasing the likelihood of separation. This lowers expected job tenure and thus the worker’s expected return from the pension, since the probability that the worker reaches tenure \( T \) and receives the pension declines. As we noted at the end of the previous subsection, this effect will reduce the maximum sustainable shirk premium, and at some point the contract breaks down. Put differently, as the likelihood of exogenous separation becomes large, the payoff date in the contract must get increasingly close to the initiation date for the worker to sustain the risk of exogenous separation. Otherwise, the worker does not expect to remain employed until \( T \), and the expected value of the pension payoff goes to zero. However, reducing \( T \) also reduces the nominal value \( W \) of the pension. At some point expected tenure \( E(\tau) \) becomes small enough that the worker will not accept the contract. Consequently, a decline in expected tenure would lead to fewer and less valuable pension contracts.

**Increased uncertainty in the productivity process.** The productivity threshold \( R \) is invariant to a change in the variance of the productivity process. Hence, a mean-preserving spread in the productivity distribution raises the probability that the match will fall below the cutoff value \( R \) at some future date, if \( R \) lies below the mean of the productivity distribution.\(^34\) Again, the terminal date \( T \) must be reduced for workers to be willing to accept the pension, but that reduces the nominal value of the pension and undermines the willingness of workers to accept the pension contract.

**Summary.** The preceding discussion provides intuition as to the breakdown of DB pensions. Matches with stable or increasing continuation values can benefit from deferring payment to the worker in order to provide incentives that are unavailable in a standard Nash bargaining model. These contracts preserve jointly efficient matches that would ordinarily be severed under Nash bargaining. Matches with decreasing continuation values, however, might not sustain the pension contract, and shifts in the stochastic process that further reduce expected future productivity of existing matches make this increasingly likely. Additionally, a mean-preserving spread in the distribution of future productivity draws may sufficiently raise uncertainty about match duration that the

\(^33\) The productivity distribution drifts stochastically in DRW.

\(^34\) The implications of a mean-preserving spread are reversed if \( R \) lies above the mean, but it is unlikely that the mean productivity draw would be sufficiently high to warrant preserving matches.
pension contract can be sustained.\textsuperscript{35}

4.5 Defined contribution pensions

The discussion above explains the purpose of DB pensions but says little about DC pensions. Deferred compensation alters the path of wages and, for workers who face borrowing constraint, the path of consumption and saving. This should reduce the appeal of pensions, according to conventional economic theory. In this light, we offer the following possible motives for the use of DC pensions:

- features such as matching rates and short vesting periods in DC pensions mimic DB pensions and encourage longer tenure;
- workers who are attracted to deferred compensation have other productive characteristics;
- workers attach some value to deferred compensation as a form of saving;
- the government establishes tax preferences for pensions because it attaches some value to deferred compensation as a form of saving.

The first two motives build on the notion of pensions as efficiency-enhancing contracts and involve at least some substitutability between DB and DC pensions. The last two motives do not exclude the use of pensions as incentive contracts but add an additional motive – to promote long-term saving. It is important to note that none of these motives, by themselves, offer an obvious explanation for the shift in pension structure and decline in job tenure.

The first explanation can be viewed as an extension of our theory of DB pensions. If pension structure shifted exogenously, then DC pensions may be used to replicate the incentives of DB pensions. The second explanation is based on the screening theory of pensions, discussed above. Earlier, we discussed reasons why we do not think that screening models fully explain the use of pensions or the shift in pension structure.

Under the last two explanations, either individuals or the government value pensions as a forced saving mechanism. The notion that individuals who otherwise have trouble saving value pensions as a commitment device arises from psychologically-based behavioral models of saving (Laibson, Repetto, and Tobacman 1998). It is a leading explanation for the findings of many researchers that both 401(k) plans and DB plans raise personal savings rates.\textsuperscript{36} Alternately, the government may establish tax preferences for retirement saving because it

\textsuperscript{35}One must consider other possible contracts at this point. Ramey and Watson (1997) showed that contracts with severance payments or punishments can sustain matches in the efficient but incentive-incompatible region. However, such contracts are rarely observed, perhaps because they are not easily enforceable or yield socially inefficient litigation upon separation.

\textsuperscript{36}Engen and Gale (2000), Poterba, Venti, and Wise (2001), and Webb (2001) are recent studies finding some effect of 401(k) plans on personal saving. Dicks-Mireaux and King (1983) and Diamond and Hausman (1984) found that DB plans raised personal saving.
believes people do not save enough. The government may think that people have trouble saving even if they do not, or that social insurance and welfare programs reduce saving through moral hazard.

5 Government Regulation in the Pension Model

In the moral hazard model, efficiency is enhanced by moving from the standard Nash contract to the DB pension contract if expected match tenure is sufficiently high. A decline in expected tenure, though, may destroy the viability of the pension contract. In this section we will evaluate the consequences if pension structure changes because of government regulation instead.

According to (12), the worker may accept the pension contract even though it involves a risk of exogenous separation. This occurs if the likelihood of a bad productivity draw is sufficiently low as to satisfy (12). Equivalently, the worker accepts the pension contract if $E(\tau)$ is sufficiently long.

Thus, some workers who take the pension contract will later suffer exogenous separation before they collect their pension. Suppose the government requires that workers be guaranteed their pension wealth regardless of whether matches reach duration $T$, even at the cost of reducing overall productivity. This destroys the firm’s ability to influence worker effort, and we can evaluate the loss resulting from rekindling the moral hazard problem. Again, suppose that output is drawn from a standard uniform distribution. Aggregate production in period $t$ under the pension contract, $\hat{Y}_P$, is the same as aggregate Nash production, which we defined earlier as $\hat{Y}_N = \int_{R_N}^{\infty} dF(y) = 1 - R_N$. Similarly, we defined aggregate production under moral hazard as $\hat{Y}_{MH} = \int_{R_{MH}}^{\infty} dF(y) = 1 - R_{MH}$. Therefore, the efficiency loss from government-imposed portability of pensions is identical to the efficiency loss $\Lambda$ from moral hazard,

$$\Lambda = \frac{R_{MH} - R_N}{1 - R_N}.$$

Figure 4 showed the increases in the efficiency loss as the shirk premium $x^w$ and the outside option $b$ rose, given other reasonable parameter values. For values of $x^w$ and $b$ around 0.5 (recall that output draws are bounded between zero and one in this example), the efficiency loss can reach 8%, while for values around 0.65 to 0.7, it can be twice as high.

Thus, our model presents the policymaker with a choice between social efficiency versus helping a fraction of workers who experience bad luck by mandating portability of pensions, which lowers aggregate productivity.
6 Trends in Pensions, Job Tenure, and Technology

In this section, we present empirical evidence that supports our hypotheses. It would be difficult to estimate our model directly, given the absence of linked employee-employer longitudinal data or even employee longitudinal data with details about the structure of compensation or productivity. The alternative is to test the model’s implications as they relate to pension structure and job tenure. While this effort is also limited by the available data, there are several types of evidence that we bring to bear.

First, we show that job tenure is related to pension structure, as our model presumes. We find that workers with DB pensions have longer tenure than workers with DC pensions or workers with no pensions. This relationship has persisted as DB pension use has declined, so the same workers are experiencing both trends. We also find that workers with more valuable DB pensions have longer tenure, and that the value of DB pensions explains much of the differential effect on tenure of having a DB versus a DC pension.

Second, we show evidence that the value of long-term jobs has dropped, which supports our explanation for the decline in job tenure and DB pensions. We find that DB pensions may have declined in value, which suggests that they play a reduced role in deterring mobility. We also show a drop in the estimated "return to tenure" – the earnings premium enjoyed by higher-tenure relative to lower-tenure workers, suggesting that the value of moving to a new job rose.

Third, we show that the decline in job tenure is related to trends in technological progress, which supports our explanation for the reason behind the decline in job tenure and DB pension use. In particular, greater declines in job tenure occurred in industries with higher rates of technological progress. This last finding supports the hypothesis that changes in the productivity process induced a shift away from long-term jobs and DB pensions; if, instead, government regulation induced the shift in pension structure, there would be no reason to expect such a link.

In the rest of this section, we use data on individuals from the Survey of Consumer Finances and the Current Population Survey. As we noted earlier, the SCF includes information on both actual and expected job tenure and on pension coverage. It began in 1983 and takes place every three years, offering the longest consistent information on pension coverage among individual-level data sets, but it has very aggregated information on industry and occupation. Therefore, we also use the CPS, which has reported consistent information on job tenure since 1983, reports three-digit industry and occupation data, and offers a very large sample. We also use information from the April 1993 CPS, the last time questions were asked about pensions.
6.1 Pension structure and job tenure

Earlier research found a link between DB pension coverage and low rates of worker mobility (for example, Allen, Clark, and McDermed 1993). While we do not estimate a model of mobility as earlier studies did, we find several interesting results. We show that workers with DB pensions have longer current and expected total job tenure than workers with DC pensions and workers without pensions, in contrast to some past research in this area. We also find that workers with more valuable DB pensions have longer tenure, controlling for the level of earnings, with the value of DB pensions explaining much of the differential effect on tenure of having a DB pension. Our approach does not allow us to distinguish whether DB pensions cause longer tenure, however.37

*Regressing job tenure on pension type.* We ran several regressions, separately for men and women, using both the SCF and CPS. Coefficient estimates are reported in Tables 4 and 5.38,39

Table 4 about here

Male workers with a DB pension in the SCF have been in their jobs about 5 years longer than workers without a pension, depending on the specification, and female workers with a DB pension have been in their jobs about 4 years longer. Workers with both a DB and DC pension have been in their job about half a year longer than workers with only a DB pension, but the difference is generally not statistically significant. In comparison, workers with a DC pension have been in their job 2-3 years longer than workers without a pension, significantly shorter than workers with a DB pension. It is unclear whether one wants to control for job characteristics such as industry and occupation, since they may help explain both pension structure and job tenure. Nevertheless, including these controls in the second and fourth columns of Table 4 reduces the estimated effect of pensions on tenure by only a year or less.

It is important to point out that the relationship between pensions and job tenure remains strong when year effects are included, so it does not reflect a spurious correlation between two trending variables. Moreover, it persists if we interact pension type with year, in results that are not shown; the estimated response to DB pensions was strong throughout the sample period and remained significantly higher than the response to DC pensions, though it declined by about 1-2 years by 1998, compared to 1983.

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37 As we noted earlier, we were unconvinced by previous attempts to estimate endogenous selection into jobs with DB pensions.
38 It would be preferable to run regressions on the probability of job exit if we had such data, rather than on current job tenure, which represents incomplete spells. The mean of complete and incomplete job spells will be the same if there is no duration dependence in spells. If spells are duration-dependent, then a linear regression on tenure can be viewed as a first-order Taylor expansion of more complicated specifications (Freeman 1980). As an alternative, we also ran regressions using total expected job tenure as the left-hand side variable, as a way to capture information about the expected duration of the spell.
39 When using SCF data, all coefficient estimates and standard errors are computed from regressions run on multiple implicates (Rubin 1987). We ran all tenure regressions separately by gender because men and women exhibited different secular trends, as shown in Tables 1 and 2. We obtain similar results if we run regressions on the log of tenure, as suggested by Even and Macpherson (1996).
Since the effect of DB pensions has been consistently strong, we can conclude that the same workers who are experiencing a decline in DB pension coverage are also spending less time in their jobs. Without ascribing a structural interpretation, we can understand the magnitude of the estimated effect by noting that the observed shift in pension structure between 1983 and 1998 is associated with a decline in job tenure of 0.9 years for males and 0.6 years for females, according to regressions (4) and (8); this is of the same order of magnitude as the observed decline.

We also ran regressions with total expected job duration as the left-hand side variable. These regressions showed a similar, significant differential effect of pension type on tenure. Workers with a DB pension have total expected tenure that is 5.5-7.5 years longer than workers without a pension, while workers with a DC pension have total expected tenure that is 3.5-4.5 years longer.

Lastly, the effect of pension structure on tenure is almost the same in the April 1993 CPS as in the SCF, as shown in Table 5. When job characteristics are included in specifications (2a), (2b), (6a), and (6b), both male and female workers with a DB pension have been in their jobs about 4 years longer than workers without a pension. In comparison, workers with a DC pension have been in their job around 2 1/2 years longer than workers without a pension, again significantly shorter than workers with a DB pension. It is also noteworthy that the estimated effect of pensions declines by only a couple tenths of a year when detailed industry and occupation codes, which are available in the CPS but not in the SCF, are included in specifications (2) and (5).

Table 5 about here

_regressing job tenure on DB pension value._ We used data reported in the SCF to compute the value of DB pensions and added that information to regressions like those reported in Table 4. Most individuals with DB pensions report the pension benefit which they expect to receive if they stay in their job as long as intended; we used this information in two different ways. In some regressions, we included this information directly. However, because it is endogenous to some degree with expected tenure and also reported with considerable error (Gustman and Steinmeier 1999), in some regressions we included instead the average expected benefit, imputed on the basis of earnings, industry, occupation, education, unionization, employer size, and gender. As in the results reported above, we control for earnings to remove spurious correlation between highly compensated jobs and long tenure.

This is similar to the approach in Gustman and Steimeier (1993, 1995), described in more detail below, of including a measure of pension backloading. Our analysis differs in two primary ways. First, their left-hand side variable is the probability of job exit and their right-hand side variable is the value of backloaded compensation that is available if the employee does not exit; a higher value of backloading is predicted to discourage exit. Our

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40The numbering of the CPS regressions in Table 5 parallels the numbering of the SCF regressions in Table 4.
left-hand side variable is tenure and our right-hand side variable is the value of the pension benefit, with a higher value predicted to encourage longer tenure. Second, we use pension data reported by the individuals in our sample (or based on information collected from the employers of the individuals in 1983); because they lacked such data, Gustman and Steinmeier imputed the backloading variable from employer-provided information in the 1983 SCF, using industry, occupation, and unionization as covariates.

We find several interesting results, which are shown in Table 6.\textsuperscript{41} First, a higher value of one’s DB pension at retirement is associated with significantly longer tenure. The results show that the semi-elasticity of tenure with respect to the monthly pension benefit (the statistic reported by most of the post-1983 sample) is 0.5-0.6 when self-reported information is included (in regressions labelled (a)) and 0.25-0.4 when the average benefit is included (in regressions labeled (b)). This implies, for a male with the median value of expected future pension benefits ($883 per month in 1998 dollars), that job tenure is 3.5 years longer than it is for someone without a DB pension, according to specification (4a), half a year longer than someone with the 25th percentile value ($351), and almost a year shorter than someone with the 75th percentile value ($1867). For a female with the median value ($541), job tenure is 3.8 years longer than it is for someone without a DB pension, according to specification (8a).\textsuperscript{42}

Second, once we control for DB pension value, then the additional effect of having a DB pension is smaller than in Table 4. It lies in a range between 0.5-3.5 years (versus 4-5 years in Table 4), in some cases not statistically distinguishable from zero and in others not statistically distinguishable from the effect of DC pensions, which remains in the range of 2-3 years. Together, these first two findings support our hypothesis that DB pensions are used to extend job tenure, since the differential effect of DB pensions on tenure is operating through the value of the pension.

Third, the semi-elasticity of tenure with respect to earnings is somewhat larger (in the range of 1.5-2.5 years) than the semi-elasticity with respect to pension value. This contradicts results in Gustman and Steinmeier (1993, 1995) which they argued were anomalous. According to their estimates, a given increase in pension backloading had a much greater effect in deterring mobility than a given increase in current compensation. We will discuss their results in greater detail next.

\textit{Comparison to other results.} The results in Tables 4-6, duplicated in two data sets, differ importantly from

\textsuperscript{41}Again, all coefficient estimates and standard errors are computed from regressions run on multiple implicates. The pension variables are described in more detail in the notes to Table 6. The numbering of the regressions in Table 6 parallels the numbering of the regressions in Table 4.

\textsuperscript{42}When we use total expected tenure on the left-hand side, the estimated semi-elasticity with respect to the same variables are about twice as large.
Gustman and Steinmeier (1993, 1995). Most notably, they found similar mobility rates for workers with DB and DC pensions. Their econometric analysis was similar, but they used different data and included proxies for variables which they had difficulty measuring. There are four key differences in their approach. First, they used data from the Survey of Income and Program Participation. The SIPP is a panel, which allowed them to focus on mobility rather than tenure. However, they only used data from 1984-85, before DC plans grew in popularity. Also, in their words, “the SIPP question sequence on plan type is atypical” (p.303, 1993) and overstated the prevalence of DC plans. The explanation for differences our pension and tenure results may therefore involve differences in the time period or in how job mobility and pension structure are measured, rather than differences in the econometric approach, which are described next.

Second, they found similar mobility rates for workers with DB and DC pensions, both in their raw data and in their econometric estimates. In our results in Tables 4 and 5, we found that workers with DC plans have shorter average tenure. Third, their next step was to replace pension type with an imputed measure of pension backloading as a independent variable, similar to our measure of pension wealth in Table 6. As we noted earlier, in both cases the estimated effect of the pension value was significant. However, in their case pension backloading had an estimated effect that was many times larger than current compensation, a finding they argued was anomalous. In our case, the effect of current earnings is of the same magnitude but is greater than the effect of pension wealth.

Fourth, their final step was to include as an explanatory variable an imputed measure of the alternative wage, adjusted for sample selection; we have not included any such variable. Their key finding was that controlling for alternative compensation reduces the estimated effect of pensions, apparently because pensioned workers face worse alternatives relative to their current job than do non-pensioned workers. In those results, the effect of pension backloading and of current compensation relative to the alternative is now of a similar magnitude, with the effect of pension backloading falling just short of statistical significance. Since they found that pension wealth is smaller than the imputed compensation differential, they concluded that the compensation differential has a greater effect in deterring mobility. This is an important finding, but it hinges on knowing the terms of alternative jobs available to workers who do not move. No motivation was offered for the identifying exclusion restrictions that Gustman and Steinmeier used in imputing alternative compensation, and we do not have a great deal of confidence that this can be done without introducing substantial bias.

To sum up, Gustman and Steinmeier did not, in our view, present convincing evidence that the alternative compensation premium, rather than the structure of DB pensions, explain the pension-mobility relationship. We

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43 The reliability of job mobility data in the SIPP is unclear, though, since it does not show the decline in job tenure which is apparent in other surveys, as we discussed earlier.

44 Age, marital status, children under 18, and home ownership were included in the mobility equation but excluded from the current and alternative compensation equations.
have reached this conclusion because of our evidence of a robust relationship between pension structure and reduced mobility, which contrasts with their evidence, and because of the difficulties of measuring alternative compensation.

6.2 The value of long-term jobs to workers

As we have emphasized, we do not have data to estimate changes in the value of a long-term job. Most importantly, we have no way to measure the employer’s share of the value of long-term jobs. Nevertheless, we can examine indicators of changes in the worker’s share. We show two ways in which tenure-related compensation appears to have shrunk: DB pensions appear to have lost value, and the degree to which earnings rise with tenure has fallen. After presenting the evidence, we discuss how these results can be interpreted in light of our model.

The value of DB pensions. The SCF does not provide ideal data to measure changes in the value of DB pensions. As described above, from 1989 on it reports the benefit people expect to receive when they leave the firm. In order to detect changes over time, we regress this variable on year dummies. However, as was clear from Figure 2, the expected pension benefit depends on two factors: the degree to which a DB pension defers compensation as well as the worker’s expected tenure, which determines how much of the deferred compensation the worker will receive. In order to isolate the first component, we include detailed controls for current and expected remaining job tenure in the regressions. We also control for current earnings, in case overall compensation declined, and in some cases we control for other individual and job characteristics, to isolate shifts in the terms of particular jobs, rather than shifts in the composition of jobs.

Table 7 about here

With the caveat that it may reflect changes in tenure to some extent, the regressions in Table 7 show that DB pensions declined significantly in value between 1989 and 1998. For the average male with a DB pension, according to the results in specification (2), the expected monthly benefit declined by $46 between 1989 and 1992, by $179 between 1992 and 1995, and by $69 between 1995 and 1998. The overall drop of $295 represented an 18% decline below the average monthly benefit in 1989 of $1648 (measured in 1998 dollars) and is statistically significant at the 90% confidence level. Among females, according to the results in (4), the trend tended to be negative but not monotonic, since the expected monthly benefit fell $322 in 1992, rose $256 in 1995, and then fell $191 in 1998. The overall decline of $258, or 26% below the 1989 value of $998, is statistically significant at the 95% level. These results suggest that deferred compensation is smaller than it used to be, especially since pension wealth in DC plans is largely independent of tenure (Gustman and Steinmeier 1995).45

45 We also verified that the decline in the value of DB pensions, if it occurred for exogenous reasons, is not nearly large enough to
The relationship between tenure and earnings. A fairly common practice in the labor literature is to estimate a "return to tenure", with current earnings on the left-hand side and tenure, along with other measures of human capital, on the right. Farber (1999) described the problem of interpreting such estimates, since many theories (including ours) predict that compensation is structured to influence tenure. While our model does not feature a rising wage profile, we alluded earlier to possible extensions in which rising wages would serve, like pensions, as an element of an incentive contract designed to lengthen tenure, deter moral hazard, encourage job-specific investment, etc. Farber argued that such estimates are interesting, nonetheless, in revealing the nature of firm-level compensation structures. This motivates our analysis as well.

For our analysis, we used data from CPSs between 1983 and 2000, since the CPS offers large sample sizes and detailed industry and occupation codes. We estimated log earnings equations for men and women separately, and we included linear and higher-order terms (up to quartics) in both tenure and potential labor market experience for each CPS year. Since we control for experience, we do not have to worry that changes in the earnings-experience relationship will bias estimates of the earnings-tenure relationship.

Table 8 shows the earnings premium paid to the average male and female worker with 5, 10, 15, 20, and 25 years of tenure, compared to a worker beginning a job. The results offer some, though not complete, support for our claim that the value of long-term jobs declined over the last twenty years. There was a sharp decline in the tenure premium, but it did not begin until the early 1990s for men and the middle 1990s for women. Before that, the tenure premium rose, though not significantly. For example, the premium to earnings enjoyed by males with 10 years of tenure rose from 20.4% in 1983 to 24.7% in 1991 and then fell to 16.5% in 2000. The overall drop in the tenure premium between 1983 and 2000 was statistically significant, and it fell the most for males with 10-15 years of tenure. For females, the drop-off occurred a little later but was sharper, so that the premium at each year of tenure shown in Table 8 was significantly lower in 2000 than in 1983. For example, for females with 10 years of tenure, the earnings premium rose from 25.5% in 1983 to 28.5% in 1996 and then fell to 14.6% in 2000.

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We did not use earlier CPS tenure supplements because the wording of the tenure question changed. We adjusted the reported tenure data for the problem, identified by Diebold, Neumark, and Polsky (1996, 1997) of half-year rounding among those with 1-2 years of tenure; while they estimated the probability of rounding up at between 56 and 65%, depending on the sample year, we simply assumed that half of individuals rounded up. We did not adjust the data for heaping at five-year intervals, as they did; in their 1996 paper, they showed that their adjustments for rounding and heaping did not affect conclusions about the magnitude of trends in job tenure. When we tried adjusting the sampling weights, as they did, for differences in nonresponse to the tenure question by age, sex, and race, the results were virtually identical. More information on these regressions is reported in the notes to Table 8.

The estimated “return to experience” fell by roughly ten percentage points for men and rose by about the same amount for women.

The F-statistic testing whether tenure was jointly significantly different in 2000 than in 1983 at all of the five-year age intervals shown in Table 8 had a p-value above 97%. As a point of reference, Topel (1991) estimated that the earnings premium for males with 10 years of tenure was over 25%.
We also find that the tenure premium dropped in industries with declines in average tenure. We determined this by regressing the median real tenure premium on average job tenure, although we do not ascribe a causal interpretation to the results. We find that a one-year decline in average job tenure is associated with a significant 2.2 percentage point decline in the tenure premium (which has a median value of roughly 18%).

**Interpretation.** Our results show that both DB pensions and the earnings premium associated with longer tenure have shrunk in value. Under what assumptions can we infer from this evidence that the value of long-term jobs has fallen? We control for the level of current earnings, so these changes are not a consequence of an overall reduction or redistribution of match surplus. In the context of our model, the evidence demonstrates that something related specifically to the duration of the job changed.

Our discussion of the pension contract illustrated the tradeoff between the term and the value of the DB pension – the worker demands a pension that pays off sooner if the likelihood of exogenous separation increases, but that reduces the size of the pension that the firm is willing to offer and at the limit threatens its viability. Thus, a decline in the value of remaining DB pensions, together with the declining use of DB pensions, supports our hypothesis that the value of long-term jobs fell. If, in contrast, increasing regulation explained the shift away from DB pensions, it is not clear why it would also reduce the value of remaining DB pensions.

The literature offers a number of explanations for observing a tenure premium. In the context of those explanations, the inferences we can draw when the tenure premium declines are generally consistent with the implications of our model. Most obviously, if the tenure premium is a component of a tenure-based incentive contract, then a decline implies that the motivation to use long-term contracts has diminished. Another possibility is that selective mobility explains the tenure premium, so that the observed wage rises with tenure because workers in better jobs stay in them longer. Selective mobility is an outcome of our matching model – matches end selectively when their productivity draw falls below a reservation level – and it will result in an observed tenure premium when it is combined with Nash bargaining and upward drift in stochastic match productivity. A shift in the productivity process that undermines the value of existing matches relative to new matches would lead to more mobility and an ambiguous effect on the observed tenure premium; those matches that persist may have even higher value, but new matches have higher values as well, so this could be consistent with our main argument. A final explanation for observing a tenure premium is that workers are paid their marginal product and that match-specific productivity rises with tenure. A decline in the observed tenure premium in this framework would result from a decline in the productivity of long-tenured workers. In sum, under various models of wage formation a decline in the observed tenure premium can be explained by a decline in the productivity

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49 We computed the average tenure premium and job tenure in 27 two-digit industries that were big enough to employ at least 100 people in each CPS survey. The regressions are weighted by the number of people from the CPS in each cell. If we control for year and industry effects (accounting for economy-wide changes and industry-specific values of the tenure premium), then the correlation is greater, with a one-year decline in average job tenure is associated with a 4.6 percentage point decline in the tenure premium.
of long-term matches. Moreover, even if the observed tenure premium is not itself a component of an incentive contract, the implied drop in the value of long-term matches can explain the move away from DB pensions.

6.3 Job tenure and new technologies

We have showed that both actual and expected job tenure fell, and that deferred compensation appears to have shrunk. From this, we infer that the value of long-term jobs has declined. In our model, this will occur if there is an acceleration of shocks that erode the productivity of existing matches relative to new matches, or even if there is simply an increase in uncertainty about the future productivity of matches. We hypothesize that a shift in the nature of new technologies, beginning 20-30 years ago, has had such effects. As above, we do not have data to test this directly, but we are able to demonstrate that the decline in job tenure in the CPS occurred in industries that experienced higher rates of technological change.

In the concluding section, we allude to a body of research finding that new technologies, and especially information-related technologies, have had the effects that we hypothesize. These earlier researchers used a variety of strategies to confront the same difficulties we face in determining the impact of new technologies. For example, Autor, Katz, and Krueger (1998) regressed an industry-level average outcome on industry-level measures of technological change, as we do.

We regressed job tenure by industry on TFP growth by industry. The earliest consistent tenure information in the CPS is reported in January 1983 and the latest in February 2000, so we computed the change in average job tenure by industry from 1983 to 2000. We use TFP growth from the Jorgenson Total Factor Productivity Series. The Jorgenson series reports annual TFP growth for 21 manufacturing sectors, along with 14 other non-manufacturing sectors of the economy from 1959 to through 1996. While it is an arbitrary choice, we used TFP growth over the ten-year period ending in 1996.50

Thus, we regressed the change in average job tenure by industry over 1983-2000 on average TFP growth by industry, weighted by the quantity of labor by industry in 1996.51 The coefficient estimate on the ten-year average of TFP growth is -32.4 (14.6), which is significant at the 95% confidence level. A one-standard deviation change in average TFP growth was 0.0111 across industries in 1996 (the range was -0.0148 to 0.0416). The estimates thus suggest that an industry with TFP growth that exceeded the average by one standard deviation experienced a decrease in average job tenure of 0.36 years; recall that average job tenure declined 0.92 years between 1983 and 1998. These correlations support our view that the decline in job tenure is linked to technological change.

50 We obtained very similar estimates using TFP growth over the five-year period ending in 1996. Additional information about the data is available in Jorgenson, Gollop, and Fraumeni (1987) and at: http://post.economics.harvard.edu/faculty/jorgenson/data/35klem.html.

51 The quantity of labor input is computed as the value of labor inputs divided by the price of labor inputs.
7 Conclusion

In this paper we have specified a model of DB pensions and job tenure. DB pensions eliminate inefficient job destruction resulting from moral hazard. In related papers, DB pensions or other forms of deferred compensation encourage investment in job-specific capital (Engemann, Friedberg, and Owyang 2003) or discourage on-the-job search (Burdett and Coles 2003; Stevens forthcoming; Friedberg, Owyang, and Sinclair 2003). The use of DB pensions is undermined, however, if expected job tenure declines. We have shown in this paper that both actual and expected job tenure fell along with the use of DB pensions.

We also used the model to demonstrate the types of changes in the stochastic productivity process which will reduce expected job tenure and hence the use of DB pensions. In particular, we focused on shocks that increase uncertainty about future match productivity, perhaps by affecting the value of existing job-specific capital. Lastly, we showed that industries which experienced more rapid growth in technological progress experienced greater declines in job tenure.

The occurrence of shocks of the type we have in mind is suggested by a large body of research on the shifting nature and pace of technological change. Researchers have found that the diffusion of new, especially information-related, technologies has had a powerful effect on the level of compensation by increasing earnings inequality.52 Other research shows rising inequality in pension and health insurance coverage as well. Our hypothesis is that it has altered the structure as well as the level of compensation by reducing the value of long-term jobs.

The key reason for rising inequality is that new technologies are largely skill and ability-biased – that is, requiring workers of greater ability, and replacing tasks performed by unskilled workers. Autor, Katz, and Krueger (1998) showed that the average skill level of workers rose more in industries that experienced higher rates of investment in general and of computerization in particular. Furthermore, case study evidence suggests that what makes new technologies ability-biased is that they require new, and in some cases more complex, skills. Computer use is, obviously, one of the new skill requirements. Both employers and individuals continue to devote substantial resources to computer training, even while computers have grown easier to use over time.53 Besides that, computers have automated routine tasks, often involved in clerical and assembly line jobs, while altering and often making more complex the performance of non-routine tasks (Levy and Murnane 1996, Autor, Levy, and Murnane 2002). Computerization often brings on further changes in required skills, workplace organization, and the delivery of services, requiring substantial training and other adjustment costs (Bresnahan, Brynjolfson, and Hitt 2002). Our evidence suggests, moreover, that jobs have been reorganized in ways that reduce the ties of workers and firms in long-term relationships.


53 For example, the University of Virginia provided computer training to 2000-3000 staff in over two thousand total workshops per year during 1998-2001. The University furnished 3.86 training hours per employee in 2001, up from 0.73 in 1994 (Friedberg 2003).
It is less clear whether these changes originated with the information revolution in the 1970s (Greenwood and Yorukoglu 1997, Caselli 1999) or started decades or a century earlier and accelerated recently (Goldin and Katz 1998, Acemoglu 2002). It is also uncertain whether recent changes represent a one-time shock or a permanent change in the stochastic productivity process. Our results provide some evidence in this regard. If agents recognize a one-time shift, then we would see them reshuffle into new matches (with some workers willing to leave jobs with DB pensions) with change in expected remaining job tenure or in the use of DB pensions in new jobs. Since both have declined, it appears that agents believe that the productivity process has permanently changed.

8 Appendix

The following is a list of variables that are used in the model:

\[ y \equiv \text{current period output (a random variable)} \]

\[ F(y) \equiv \text{cumulative distribution function of output} \]

\[ Y_{t,t} = Y \equiv \text{realized current period output} \]

\[ g(R) = g_f + g_w \equiv \text{match continuation value for firm (f), worker (w)} \]

\[ \phi = \phi_f + \phi_w \equiv \text{value of reentering matching pool} \]

\[ R_N \equiv \text{reservation level of output } y \text{ under pure Nash bargaining} \]

\[ R_{MH} \equiv \text{reservation level of output } y \text{ under Moral Hazard} \]

\[ b = b_f + b_w \equiv \text{contemporaneous outside option for firm (f), worker (w)} \]

\[ \lambda \equiv \text{probability of finding a match when one is searching} \]

\[ \beta \equiv \text{discount rate, } 0 < \beta < 1 \]

\[ \theta \equiv \text{worker’s relative bargaining power, } 0 < \theta < 1 \]

\[ x^w \equiv \text{worker’s payoff to shirking} \]

\[ w \equiv \text{current period wage} \]

\[ W \equiv \text{pension payoff at tenure } T \]
References


### Table 1: Current job tenure

<table>
<thead>
<tr>
<th></th>
<th>Men Average, by years of potential experience</th>
<th>Women Average, by years of potential experience</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>0-5</td>
<td>6-15</td>
</tr>
<tr>
<td>1983</td>
<td>2.6</td>
<td>4.9</td>
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<td>1989</td>
<td>2.0</td>
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<tr>
<td>1992</td>
<td>2.2</td>
<td>4.8</td>
</tr>
<tr>
<td>1995</td>
<td>1.8</td>
<td>4.5</td>
</tr>
<tr>
<td>1998</td>
<td>1.6</td>
<td>4.4</td>
</tr>
<tr>
<td>change,</td>
<td>-1.0</td>
<td>-0.5</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>HS or less Average, by education and years of potential experience</th>
<th>Some college Average, by education and years of potential experience</th>
</tr>
</thead>
<tbody>
<tr>
<td>1983</td>
<td>4.3</td>
<td>11.6</td>
</tr>
<tr>
<td>1989</td>
<td>3.8</td>
<td>11.1</td>
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<tr>
<td>1992</td>
<td>4.4</td>
<td>10.2</td>
</tr>
<tr>
<td>1995</td>
<td>3.6</td>
<td>10.0</td>
</tr>
<tr>
<td>1998</td>
<td>3.9</td>
<td>10.1</td>
</tr>
<tr>
<td>change,</td>
<td>-0.4</td>
<td>-1.5</td>
</tr>
</tbody>
</table>

Data source: Survey of Consumer Finances from 1983, 89, 92, 95, 98. Respondents were asked “How many years in total have you worked for this employer?” Sample: Full-time employees, except those who reported tenure that exceeded potential experience plus two years (about 1.5% of the sample). Details: Statistics were weighted using survey weights. Years of potential experience is defined as age minus years of completed education minus six.
## Table 2: Expected remaining job tenure

<table>
<thead>
<tr>
<th></th>
<th>Men Average, by years of current tenure</th>
<th>Women Average, by years of current tenure</th>
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</thead>
<tbody>
<tr>
<td></td>
<td>0-5</td>
<td>6-15</td>
</tr>
<tr>
<td>1983</td>
<td>19.6</td>
<td>20.6</td>
</tr>
<tr>
<td>1989</td>
<td>16.7</td>
<td>17.3</td>
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<tr>
<td>1992</td>
<td>17.5</td>
<td>19.0</td>
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<tr>
<td>1995</td>
<td>15.0</td>
<td>17.2</td>
</tr>
<tr>
<td>1998</td>
<td>14.5</td>
<td>16.9</td>
</tr>
<tr>
<td>change, 1983-98</td>
<td>-5.2</td>
<td>-3.6</td>
</tr>
<tr>
<td>change, 1989-98</td>
<td>-2.2</td>
<td>-0.3</td>
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</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>HS or less Average, by education and years of current tenure</th>
<th>Some college Average, by education and years of current tenure</th>
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</thead>
<tbody>
<tr>
<td>1983</td>
<td>21.8</td>
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<tr>
<td>1989</td>
<td>17.8</td>
<td>14.4</td>
</tr>
<tr>
<td>1992</td>
<td>20.7</td>
<td>12.7</td>
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<tr>
<td>1995</td>
<td>18.3</td>
<td>12.8</td>
</tr>
<tr>
<td>1998</td>
<td>18.1</td>
<td>13.6</td>
</tr>
<tr>
<td>change, 1983-98</td>
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<td>0.0</td>
</tr>
<tr>
<td>change, 1989-98</td>
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<td>-0.8</td>
</tr>
</tbody>
</table>

Data source: Respondents were asked “How many years do you expect to continue working for this employer?”, Survey of Consumer Finances from 1983, 89, 92, 95, 98.
Sample: full-time employees aged 21-59
Details: Statistics were weighted using survey weights. Approximately 14% of respondents answered that they would “never stop”; we took the following steps to impute a specific answer for them: (1) we used their answer if they responded to a later question about when they would retire from all work; or else (2) we used their answer if they responded to a later question about when they would retire from full-time work; or else (3) we assumed that they would work until the age of eighty.
In 1995-98, “less than a year” was coded as a separate answer, in which case we assigned a value of zero; in 1983-92 one is the smallest coded response, and for respondents who were coded with a value of one, we randomly assigned an answer of zero in the same proportion as is observed among those answering zero or one in 1995-98 (which will lead to a slight underestimate of the decline in tenure).
### Table 3: Analysis of variance in pension coverage

<table>
<thead>
<tr>
<th>variation in pension coverage explained by year effects:</th>
<th>Dependent variable</th>
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<tbody>
<tr>
<td></td>
<td>Has a:</td>
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<tr>
<td></td>
<td>pension</td>
</tr>
<tr>
<td>all year effects (fx)</td>
<td>100%</td>
</tr>
<tr>
<td>year main fx</td>
<td>12.1</td>
</tr>
<tr>
<td>year*education fx</td>
<td>16.7</td>
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<tr>
<td>year*occupation fx</td>
<td>26.7</td>
</tr>
<tr>
<td>year*industry fx</td>
<td>35.1</td>
</tr>
<tr>
<td>year*union fx</td>
<td>9.4</td>
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</table>

Data source: Survey of Consumer Finances from 1983, 89, 92, 95, 98.
Sample: Full-time employees.
Details: The columns of this table report results from ANOVA estimates; each column reports results with a different dependent variable. In addition to the year effects and interactions shown in the table, the independent variables include age and age squared; employer size; main effects for education (4 categories), occupation (6 categories), industry (6 categories) and union coverage; interactions of education and occupation and of industry and occupation; gender and interactions of gender with education, occupation, and industry.
Statistics were weighted using survey weights. The surveys from 1989 and later include five sets of multiply imputed variables (called implicates). While other estimates involving variance in this paper correct for the presence of multiple imputations using the methods introduced in Rubin (1987), the estimates in this table do not; rather, the estimates treat each implicate as independent data. However, ANOVA estimates done on each implicate separately did not deviate by more than two percentage points in either direction from those reported here.
Table 4: Job tenure and pension coverage (OLS regression results, SCF)

Dependent variable: years of current job tenure

<table>
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<tr>
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<th>(3)</th>
<th>(4)</th>
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</thead>
<tbody>
<tr>
<td>Men</td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(mean of dependent variable = 8.77)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>has DB pension only</td>
<td>5.49*** (0.27)</td>
<td>4.24*** (0.29)</td>
<td>5.47*** (0.28)</td>
<td>4.44*** (0.30)</td>
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<tr>
<td>has DC pension only</td>
<td>2.54*** (0.28)</td>
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<td>2.55*** (0.28)</td>
<td>2.31*** (0.28)</td>
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<tr>
<td>has DB &amp; DC pension</td>
<td>5.69*** (0.35)</td>
<td>5.00*** (0.34)</td>
<td>5.66*** (0.35)</td>
<td>5.05*** (0.35)</td>
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<table>
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<td>Women</td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(mean of dependent variable = 7.62)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>has DB pension only</td>
<td>4.07*** (0.35)</td>
<td>3.45*** (0.36)</td>
<td>4.08*** (0.36)</td>
<td>3.56*** (0.37)</td>
</tr>
<tr>
<td>has DC pension only</td>
<td>2.25*** (0.29)</td>
<td>2.10*** (0.28)</td>
<td>2.21*** (0.28)</td>
<td>2.09 (0.27)</td>
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<tr>
<td>has DB &amp; DC pension</td>
<td>4.38*** (0.38)</td>
<td>3.85*** (0.40)</td>
<td>4.43*** (0.39)</td>
<td>3.96*** (0.40)</td>
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Regression also includes:

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<td>yes</td>
<td>yes</td>
<td>yes</td>
</tr>
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<td>job variables</td>
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<td>no</td>
<td>yes</td>
</tr>
<tr>
<td>year effects</td>
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<td>no</td>
<td>yes</td>
<td>yes</td>
</tr>
<tr>
<td>year*job variables</td>
<td>no</td>
<td>no</td>
<td>no</td>
<td>yes</td>
</tr>
</tbody>
</table>

Data source: Survey of Consumer Finances 1983, 89, 92, 95, 98.
Sample: Full-time employees, excluding those who report tenure in excess of potential experience plus two (about 1.5% of the sample); those whose pension type is unknown (approximately 0.5% of the remaining sample); and those with earnings in the top or bottom 1% of the distribution.
Details: The coefficient estimates and Huber-White standard errors are computed from regressions run on multiple implicates, as in Rubin (1987). The regressions were weighted using survey weights. * indicates a confidence level of at least 90%, ** 95%, *** 99%.
Specifications: (1) and (5) includes real weekly earnings (in 1998 dollars), age and age squared. (2) and (6) add job variables (4 education, 6 industry, 6 occupation, and 6 firm size dummies, industry* occupation, education*occupation, union coverage). (3) and (7) add year dummies. (4) and (8) add variables from (2) and (3) along with year*industry, year*occupation, year*education, year*union coverage.
Table 5: Job tenure and pension coverage (OLS regression results, 1993 CPS)

<table>
<thead>
<tr>
<th>Dependent variable: years of current job tenure</th>
<th>Men (mean of dependent variable = 10.0)</th>
<th>Women (mean of dependent variable = 8.37)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td>(2a)</td>
<td>(2b)</td>
</tr>
<tr>
<td>(5)</td>
<td>(6a)</td>
<td>(6b)</td>
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</table>

Independent variables:

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<th>(2b)</th>
<th>(5)</th>
<th>(6a)</th>
<th>(6b)</th>
</tr>
</thead>
<tbody>
<tr>
<td>has DB pension only</td>
<td>5.04*** (0.33)</td>
<td>3.89*** (0.33)</td>
<td>4.11*** (0.33)</td>
<td>4.41*** (0.27)</td>
<td>3.61*** (0.28)</td>
<td>3.82*** (0.28)</td>
</tr>
<tr>
<td>has DC pension only</td>
<td>2.98*** (0.30)</td>
<td>2.65*** (0.29)</td>
<td>2.83*** (0.29)</td>
<td>2.81*** (0.26)</td>
<td>2.52*** (0.26)</td>
<td>2.65*** (0.27)</td>
</tr>
<tr>
<td>has DB &amp; DC pension</td>
<td>5.22*** (0.34)</td>
<td>4.29*** (0.34)</td>
<td>4.57*** (0.34)</td>
<td>4.98*** (0.30)</td>
<td>4.30*** (0.30)</td>
<td>4.53*** (0.30)</td>
</tr>
</tbody>
</table>

Regression also includes:

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2a)</th>
<th>(2b)</th>
<th>(5)</th>
<th>(6a)</th>
<th>(6b)</th>
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<td>age</td>
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<td></td>
<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>job variables</td>
<td>no</td>
<td>yes</td>
<td></td>
<td></td>
<td>no</td>
<td></td>
</tr>
<tr>
<td>SCF job variables</td>
<td>no</td>
<td>no</td>
<td></td>
<td></td>
<td>yes</td>
<td></td>
</tr>
</tbody>
</table>

Sample: Full-time employees who were working or had a job but were not at work, excluding those who report tenure in excess of potential experience plus three: those whose pension type is unknown (approximately 0.5% of the sample); and those with earnings in the top or bottom 1% of the distribution. The sample size is 5,830 in (1), (2), (3), and 4,892 in (4), (5), and (6).
Details: The numbering of the regressions parallels the numbering in Table 4. Weighted using survey weights. Huber-White standard errors appear in parentheses; ** indicates a confidence level of at least 90%, *** 95%, **** 99%.
Specifications: (1) and (4) includes the natural log of real weekly earnings, age and age squared. (2), (3), (5) and (6) add job variables; each includes 4 education dummies, 8 firm size dummies, and union coverage; (2) and (5), exploiting the richer data available in the CPS, include 51 industry dummies and 45 occupation dummies; (3) and (6), designed to replicate the SCF regressions reported in Table 4, include 6 industry and 6 occupation dummies, industry* occupation, and education*occupation.
Table 6: Job tenure and DB pension characteristics (OLS regression results, SCF)

<table>
<thead>
<tr>
<th>Dependent variable: years of current job tenure</th>
<th>Men (mean of dependent variable = 8.77)</th>
<th>(1a)</th>
<th>(4a)</th>
<th>(1b)</th>
<th>(4b)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Independent variables:</td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>has DB pension only</td>
<td>2.41*** (0.36)</td>
<td>1.34*** (0.47)</td>
<td>3.31*** (0.34)</td>
<td>2.72*** (0.46)</td>
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</tr>
<tr>
<td>has DC pension only</td>
<td>2.69*** (0.27)</td>
<td>2.48*** (0.28)</td>
<td>2.64*** (0.28)</td>
<td>2.36*** (0.28)</td>
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<tr>
<td>has DB &amp; DC pension</td>
<td>2.49*** (0.44)</td>
<td>1.84*** (0.52)</td>
<td>3.40*** (0.42)</td>
<td>3.24*** (0.52)</td>
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</tr>
<tr>
<td>DB pension benefits at retirement (natural log of real present value, 1998 dollars):</td>
<td>Individual-reported</td>
<td>Average</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>log value of monthly benefit</td>
<td>0.52*** (0.06)</td>
<td>0.52*** (0.08)</td>
<td>0.35*** (0.05)</td>
<td>0.26*** (0.08)</td>
<td></td>
</tr>
<tr>
<td>log value of lump-sum benefit</td>
<td>0.38*** (0.13)</td>
<td>0.41*** (0.12)</td>
<td>0.29*** (0.13)</td>
<td>0.27*** (0.12)</td>
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</tr>
<tr>
<td>log value of pension wealth</td>
<td>0.54*** (0.05)</td>
<td>0.52*** (0.05)</td>
<td>0.43*** (0.05)</td>
<td>0.41*** (0.05)</td>
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</tr>
<tr>
<td>log weekly earnings</td>
<td>1.46*** (0.27)</td>
<td>2.39*** (0.28)</td>
<td>1.51*** (0.27)</td>
<td>2.45*** (0.28)</td>
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</tr>
<tr>
<td>Women (mean of dependent variable = 7.62)</td>
<td></td>
<td>(5a)</td>
<td>(8a)</td>
<td>(5b)</td>
<td>(8b)</td>
</tr>
<tr>
<td>Independent variables:</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>has DB pension only</td>
<td>1.24* (0.73)</td>
<td>0.63 (0.83)</td>
<td>2.12*** (0.47)</td>
<td>1.72*** (0.55)</td>
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<tr>
<td>has DC pension only</td>
<td>2.39*** (0.27)</td>
<td>2.29*** (0.26)</td>
<td>2.35*** (0.28)</td>
<td>2.21*** (0.27)</td>
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<tr>
<td>has DB &amp; DC pension</td>
<td>1.57*** (0.63)</td>
<td>0.95 (0.76)</td>
<td>3.47*** (0.46)</td>
<td>2.10*** (0.56)</td>
<td></td>
</tr>
<tr>
<td>DB pension benefits at retirement (natural log of real present value, 1998 dollars):</td>
<td>Individual-reported</td>
<td>Average</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>log value of monthly benefit</td>
<td>0.57*** (0.12)</td>
<td>0.60*** (0.16)</td>
<td>0.39*** (0.08)</td>
<td>0.37*** (0.10)</td>
<td></td>
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<tr>
<td>log value of lump-sum benefit</td>
<td>0.15 (0.13)</td>
<td>0.19 (0.15)</td>
<td>0.06 (0.12)</td>
<td>0.06 (0.13)</td>
<td></td>
</tr>
<tr>
<td>log value of pension wealth</td>
<td>0.33*** (0.08)</td>
<td>0.31*** (0.08)</td>
<td>0.22*** (0.07)</td>
<td>0.21*** (0.07)</td>
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</tr>
<tr>
<td>log weekly earnings</td>
<td>1.86*** (0.75)</td>
<td>2.59*** (0.35)</td>
<td>1.94*** (0.31)</td>
<td>2.65*** (0.36)</td>
<td></td>
</tr>
</tbody>
</table>

Regression also includes:
- age: yes, yes, yes, yes
- job variables: no, yes, no, yes
- year effects: no, yes, no, yes
- year*job variables: no, yes, no, yes

Details: These regressions replicate those appearing in Table 4, with the addition of variables representing DB pension benefits expected at retirement. The value was reported in one of three different ways: (1) over 95% of individuals with a DB pension in 1989-98 reported a periodic amount that they expect to receive when they leave their job; (2) about 2.5% of individuals with a DB pension in 1989-98 reported a lump-sum amount which they expect to receive; (3) the SCF reported expected pension wealth for 55% of individuals with a DB pension in 1983, based on information collected from employers. We included the natural log of the present value of each of these variables, along with dummy variables indicating which of the three variables (if any) was reported for a given observation. In regressions (1a), (4a), (5a), and (8a), the self-reported variable is included. In regressions (1b), (4b), (5b), and (8b), the average value is included, imputed on the basis of log earnings, industry, occupation, education, unionization, and employer size, separately for men and women.

The numbering of the regressions parallels the numbering in Table 4. Huber-White standard errors appear in parentheses; * indicates a confidence level of at least 90%, ** 95%, *** 99%.

For additional information, see notes to Table 4.
Table 7: Changes in the value of DB pensions (OLS regression results, SCF)

Dependent variable: expected monthly pension benefit (1998 dollars)

<table>
<thead>
<tr>
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<tbody>
<tr>
<td>Men</td>
<td></td>
<td></td>
</tr>
<tr>
<td>year dummy, 1992</td>
<td>-8 (170)</td>
<td>-46 (187)</td>
</tr>
<tr>
<td>year dummy, 1995</td>
<td>-120 (141)</td>
<td>-226 (158)</td>
</tr>
<tr>
<td>year dummy, 1998</td>
<td>-174 (164)</td>
<td>-295* (172)</td>
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<table>
<thead>
<tr>
<th></th>
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<tbody>
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<td>Women</td>
<td></td>
<td></td>
</tr>
<tr>
<td>has DB pension only</td>
<td>-294*** (119)</td>
<td>-322*** (138)</td>
</tr>
<tr>
<td>has DC pension only</td>
<td>-25 (198)</td>
<td>-66 (223)</td>
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<tr>
<td>has DB &amp; DC pension</td>
<td>-228* (131)</td>
<td>-258** (124)</td>
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Regression also includes:

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<td>yes</td>
</tr>
<tr>
<td>job variables</td>
<td>no</td>
<td>yes</td>
</tr>
</tbody>
</table>

Data source: Survey of Consumer Finances 1989, 92, 95, 98.
Sample: Full-time employees with DB pensions who report their expected monthly benefit, excluding those who report tenure in excess of potential experience plus two (about 1.5% of full-time employees); those whose pension type is unknown (approximately 0.5% of the remaining sample); those with earnings in the top or bottom 1% of the distribution; those who report that they will receive a lump-sum benefit (2.5% of the remaining sample) and those who do not report a benefit (2.5% of the remaining sample).
Details: The coefficient estimates and Huber-White standard errors are computed from regressions run on multiple implicates, as in Rubin (1987). The regressions were weighted using survey weights. * indicates a confidence level of at least 90%, ** 95%, *** 99%.
Specifications: (1) and (3) includes real weekly earnings (in 1998 dollars), age and age squared, potential experience and experience squared, current tenure (linear through quartic terms), and expected future tenure (linear through quartic terms). (2) and (4) add job variables (4 education, 6 industry, 6 occupation, and 6 firm size dummies, industry* occupation, education*occupation, union coverage).
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<td>44.3</td>
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<td>23.4</td>
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Sample: Employees aged 21-59 who were working or had a job but were not at work, excluding those who report earnings in the top or bottom 1% of the distribution. The sample size is 32,806 men and 31,124 women.
Details: Each cell in this table reports the estimated effect of years of job tenure on the natural log of the real wage, expressed as a percentage increase associated with a given number of years of tenure. These estimates are obtained from regressions run separately on men and women. The regressions include years of job tenure (with nonlinear terms included up to the fourth power), all interacted with the CPS year; years of potential experience (up to the fourth power), all interacted with the CPS year; a dummy for being a usual full-time worker; dummies for four education categories, all interacted with the CPS year; 51 industry dummies, interacted with a dummy for being in the public sector; and 45 occupation dummies, interacted with a dummy for being in the public sector. The real wage is defined as weekly earnings divided by usual weekly hours.
All of the estimated earnings premia are significantly different from zero. The asterisks next to each cell indicate the significance level on an F-test that compares the earnings premium with the one reported in the cell below. * indicates a confidence level on the F-statistics of at least 90%, ** 95%, *** 99%. These tests are based on Huber-White standard errors. Weighted using the outgoing rotation group weights.
Figure 1: Full-time employees

- Has a pension
- Pension is defined benefit
- Pension is defined contribution
Figure 2: Accrual of Pension Wealth

Accrual, $1,000

Age at which the employee leaves the job

- Typical defined benefit pension
- Typical defined contribution pension
Figure 3

The diagram illustrates a graph with axes labeled $x$, $y+g$, and $\phi$. Points marked as $x$, $b$, and $Z$ are connected by lines labeled IC.
Figure 5

Pension Payoff (as % Nash Value)

Shirk Premium, x^w

T = 30
T = 25
T = 20
T = 15